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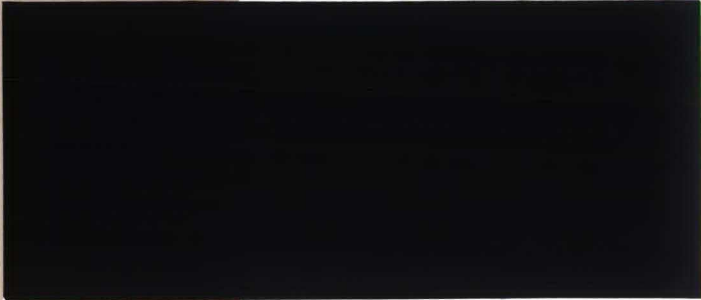No. 9006

KEYNESIAN AND NEW CLASSICAL MODELS
OF UNEMPLOYMENT REVISITED

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KEYNESIAN AND NEW CLASSICAL
MODELS OF UNEMPLOYMENT
REVISITED*

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Abstract

Several Keynesian and New Classical models of unemployment for the U.S. are re—evaluated. The basic two—equation system of the New Classical model comprises a univariate structural equation of unemployment together with a univariate expectations equation. The difference between actual and expected real federal government expenditure relative to its normal level leads to an extension of the basic New Classical model from a two—equation system to a three—equation system, namely a univariate structural equation together with a bivariate expectations system. Since estimation by two—step or multivariate two—step methods is generally neither efficient nor provides consistent estimators of the standard errors for the New Classical models of unemployment available in the literature, maximum likelihood methods are used for estimating and testing the New Classical models. Existing empirical New Classical models of unemployment are improved by expanding the set of variables used. The original and revised models are examined for adequacy by: (i) testing the cross—equation restrictions in the three—equation system; (ii) testing the significance of the anticipated and unanticipated components of monetary policy when the cross—equation restrictions are imposed; (iii) using diagnostic checks in a systems context; (iv) testing against non—nested Keynesian alternatives in both single—equation and systems contexts. The adequacy of the Keynesian model is examined by: (i) using diagnostic checks in a single—equation context; (ii) testing against the original and revised non—nested New Classical alternatives in both single—equation and systems contexts. Robustness of the outcomes of various hypothesis tests and diagnostic checks is evaluated by extending the sample period from 1946–73 to 1946–85, and these results are compared with those available in the literature. The revised New Classical model for the 1946–73 period is found to be adequate when it is estimated over the longer time period, whereas the Keynesian model is not. Moreover, it is shown that the existing results of tests obtained at the single—equation level are not always supported when the correct test statistics are calculated using single—equation estimation or when the full system of New Classical equations is estimated and tested using maximum likelihood methods.

"Let us weigh the one against the other."

> Sherlock Holmes to Dr. Watson
> in *The Adventure of the Priory School* by A. Conan Doyle

"I think that both inferences are permissible."

> Sherlock Holmes to Stanley Hopkins
> in *The Adventure of Black Peter* by A. Conan Doyle

1. Introduction

The policy ineffectiveness proposition of the New Classical school states that only unanticipated changes in the money supply affect real variables such as the unemployment rate or the level of output. At the vanguard of attempts at the empirical validation of the proposition using U.S. data was Barro (1977, 1978, 1979, 1981), with support from, among a host of others, Barro and Rush (1980), Liederman (1980), Rush (1986), and Rush and Waldo (1988). Many opponents have argued against the proposition from both empirical and methodological viewpoints, and prominent among these have been Small (1979), Mishkin (1982), Gordon (1982) and Pesaran (1982, 1988).

Although much empirical research has been undertaken for various countries using different data and different sample periods, perhaps the most revealing recent interchange has taken place between Rush and Waldo (1988) and Pesaran (1988). This debate is of interest primarily because Pesaran (1982) produced a viable non—nested Keynesian (or activist) model of unemployment which rejected Barro's (1977) model without itself being rejected by the New Classical model. Rush and Waldo (1988) argued that Pesaran's (1982) version of the New Classical model could be improved by taking account of the fact that when it is known that a war is over, the public will anticipate a reduction in government spending. They argued that the Keynesian model proposed by Pesaran (1982) could be rejected in favour of their improved New Classical model. However, Rush and Waldo's argument was easily overturned when Pesaran (1988) used the same argument to improve

the Keynesian model which, not surprisingly, was once again found to be superior to the improved New Classical model.

While the latest round in the battle seems to have been won by the Keynesian model of unemployment for the U.S., the most recent papers go beyond previous research using Barro's (1977) data in two important respects:

(i) serious attempts have been made to derive more viable non−nested alternative models of unemployment than those of Barro (1977, pp.108−109): Pesaran (1982, p.535) argues that a 'proper test' of an hypothesis "invariably requires consideration of at least one *genuine* alternative";

(ii) the Keynesian and New Classical models have been subjected to serious diagnostic tests (see Pesaran (1988)) that are a far cry from the usual provision of an adjusted coefficient of determination, a standard error of estimate and (possibly) a Durbin−Watson statistic as the mainstay of empirical research in economics.

In spite of these empirical advances, however, there are some problems that remain unresolved by the latest research efforts. In particular, the values of the anticipated and unanticipated variables present in the New Classical models are typically unobserved, and hence are generated as the predicted values and the residuals, respectively, from an auxiliary regression. Interest in such models centres on the consistency and efficiency of ordinary least squares/two step estimators (OLS/2SE), as well as consistent estimation of standard errors for valid inferences to be made. Although Pesaran (1988, footnote 2) notes that the 2SE standard errors of the New Classical model of unemployment suffer from the "generated regressors" problem analysed by Pagan (1984, 1986), no mention is made of the inefficiency of 2SE for the same problem (see McAleer and McKenzie (1988) for very simple alternative proofs of several of Pagan's efficiency results). Moreover, several of the diagnostic and non−nested tests based on 2SE also suffer from the problem of inconsistent standard errors, so that the resulting inferences might need to be re−examined. Fortunately, Theorem 8 of Pagan (1984) can be used to show that the diagnostic and

non—nested tests based on the procedure of variable addition and estimated by two step methods have calculated statistics that are, in general, biased towards rejection of the relevant null hypotheses; an identical result has also been presented in Theorem 1 of Murphy and Topel (1985), although the authors assume, rather than prove, that the error variance is estimated consistently. Thus, non—rejection of a null is a valid inference since the decision cannot be overturned using the correct statistic, whereas rejection of a null needs to be re—evaluated. Such a re—evaluation in the context of multivariate two—step estimators (M2SE) is one of the purposes of the present paper.

Although the use of diagnostic and non—nested tests has been encouraged in recent years (see, for example, Kramer et al. (1985) and McAleer et al. (1985)), there are alternative ways of testing the validity of models in a systems framework. In the context of the New Classical system, in particular, it is possible to test for the statistical significance of the anticipated and unanticipated components of monetary policy, as well as to test the cross—equation restrictions arising from the structure of the system. The New Classical model of Rush and Waldo (1988) can also be improved using existing variables. It is not necessary to look far and wide, especially since it turns out that one of the best available New Classical models is to be found in Pesaran (1982). Indeed, Pesaran's New Classical model can be shown to be superior to that of Rush and Waldo (1988), and also provides a more serious contender to Pesaran's Keynesian model of unemployment.

The purpose of this paper is to re—evaluate the existing Keynesian and New Classical models of unemployment for the U.S.. The basic two equation system of the New Classical model comprises a univariate structural equation of unemployment together with a univariate expectations equation. The difference between actual and expected real federal government expenditure relative to its normal level leads to an extension of the New Classical model from a two—equation system to a three—equation system, namely a univariate structural equation together with a bivariate expectations system. Since estimation by two—step or multivariate two—step methods is generally neither efficient nor

provides consistent estimators of the standard errors for the New Classical models of unemployment available in the literature, maximum likelihood methods are used for estimating and testing the New Classical models. The existing empirical New Classical models of unemployment are improved by expanding the set of variables used. The original and revised models are examined for adequacy by: (i) testing the cross–equation restrictions in the three–equation system; (ii) testing the significance of the anticipated and unanticipated components of monetary policy when the cross–equation restrictions are imposed; (iii) using diagnostic checks in a systems context; (iv) testing against non–nested Keynesian alternatives in both single–equation and systems contexts. The adequacy of the Keynesian model is examined by: (i) using diagnostic checks in a single–equation context; (ii) testing against the original and revised non–nested New Classical alternatives in both single–equation and systems contexts. Robustness of the outcomes of various hypothesis tests and diagnostic checks is evaluated by extending the sample period from 1946–73 to 1946–85, and these results are compared with those available in the literature. The revised New Classical model for the 1946–73 period is found to be adequate when it is estimated over the longer time period, whereas the Keynesian model is not (as shown in Pesaran (1988)). Moreover, it is shown that the existing results of tests obtained at the single–equation level are not always supported when the correct test statistics are calculated using single–equation estimation or when the full system of New Classical equations is estimated and tested using maximum likelihood methods.

The plan of the paper is as follows. In Section 2 the variables are defined and the model specifications are given. The data and sample periods used are discussed in Section 3, and the bias of some diagnostic and non–nested tests based on the variable addition method in the context of 2SE and M2SE of New Classical models is analysed in Section 4. Empirical results are given in Section 5 and some concluding remarks in Section 6.

2. Model Specifications

The original and revised Keynesian and New Classical models are given as follows:

Original Keynesian model: Pesaran (1988, equation (1), 1946–73)

$$UN_t = \phi_0 + \phi_1 MIL_t + \phi_2 MINW_t + \phi_3 DM_t + \phi_4 DM_{t-1}$$
$$+ \phi_5 DG_t + \phi_6 t + \phi_7 WAR_t + error_t \tag{1}$$

Revised Keynesian model: Pesaran (1988, Appendix Table 2, 1946–85)

$$UN_t = \psi_0 + \psi_1 MIL_t + \psi_2 UN_{t-1} + \psi_3 DM_t + \psi_4 DM_{t-1}$$
$$+ \psi_5 DM_{t-2} + \psi_6 t + \psi_7 WAR_t + error_t \tag{2}$$

Original New Classical model: Barro (1977), Pesaran (1982, 1988), Rush and Waldo (1988)

$$UN_t = \alpha_0 + \alpha_1 MIL_t + \alpha_2 MINW_t + \alpha_3 DMRH_t + \alpha_4 DMRH_{t-1}$$
$$+ \alpha_5 DMRH_{t-2} + error_t \tag{3}$$

where $DMRH_t = DM_t - E_{t-1}(DM_t)$ is the error term in the money supply equation given by

$$DM_t = \beta_0 + \beta_1 DM_{t-1} + \beta_2 DM_{t-2} + \beta_3 UN_{t-1}$$
$$+ \beta_4 E_{t-1}(FEDV_t) + DMRH_t \tag{4}$$

where $E_{t-1}(FEDV_t) = FEDV_t - 0.8 DGR_t$ and $DGR_t = DG_t - E_{t-1}(DG_t)$ is the error term in the government expenditure equation given by

$$DG_t = \gamma_0 + \gamma_1 DG_{t-1} + \gamma_2 UN_{t-1} + \gamma_3 WAR_t + DGR_t \tag{5}$$

Revised New Classical model: Pesaran (1982, Table 5)

$$UN_t = \alpha_0 + \alpha_1 MIL_t + \alpha_2 MINW_t + \alpha_3 DMRH_t + \alpha_4 DMRH_{t-1}$$
$$+ \alpha_5 DMRH_{t-2} + \alpha_6 DGR_{t-1} + \alpha_7 t + error_t \tag{6}$$

together with equations (4) and (5).

The variables are defined as follows:

$UN_t = \log[U_t/(1-U_t)]$

U_t = annual average unemployment rate

MIL_t = measure of military conscription

$MINW_t$ = minimum wage variable

DM_t = rate of growth of money supply (M1 definition)

$DMRH_t = DM_t - E_{t-1}(DM_t)$ = unanticipated rate of growth of money supply

$FEDV_t$ = real federal government expenditure relative to its normal level

$E_{t-1}(FEDV_t)$ = anticipated value of $FEDV_t$ formed at time t−1

DG_t = rate of growth of real federal government expenditure

$DGR_t = DG_t - E_{t-1}(DG_t)$ = unanticipated rate of growth of real federal government expenditure

WAR_t = a dummy variable measuring the intensities of different wars

t = time trend.

Although we are principally interested in explaining the unemployment rate because it is the focus of the debate between the competing Keynesian and New Classical models, the money and government expenditure growth rates are needed to obtain estimates of the monetary and fiscal shocks. Specifically, the money growth equation is used to obtain systems estimates of anticipated monetary policy and unanticipated monetary shocks. The government expenditure growth equation is used to obtain the systems estimates of the government expenditure shock in order to generate the expected value of real federal government expenditure relative to its normal value, since the market is not likely to be able to anticipate the current fiscal policy variable perfectly (see Mishkin (1982, p.42) and Pesaran (1982, p.540)). In specifying the government expenditure equation, it is implicitly assumed that the value of WAR_t is known to economic agents at time t−1, that is, WAR_t is perfectly predictable at time t−1. Barro (1977) specifies the rate of growth of the money

supply as a function of its own past, a measure of lagged unemployment to capture countercyclical monetary policy, and a current fiscal policy variable to account for government financing needs. The rate of growth of government expenditure, which is used to obtain the current anticipated fiscal policy variable, includes its own lag to capture the effects of any persistence in fiscal growth, a lagged value of unemployment to measure countercyclical fiscal policy, and a dummy variable for war since the public will anticipate an abrupt reduction in government military spending when a war ends (see Pesaran (1988) and Rush and Waldo (1988)). Finally, the New Classical unemployment equation is postulated to depend upon current and lagged monetary shocks and two real variables to explain the natural rate of unemployment, namely a measure of military conscription and a minimum wage variable. Barro (1977, p.107) argues that the effects of a selective military draft would tend to lower the unemployment rate, while the impact of the minimum wage rate could affect unemployment positively or negatively.

The non—nested Keynesian (or activist) reduced form alternative model developed in Pesaran (1982, 1988) takes account of the same military conscription, minimum wage and war variables as specified in the New Classical model, together with the rates of growth of the money supply and real federal government expenditure, and a time trend to explain gradual changes in the natural rate of unemployment over time. The revised Keynesian model incorporates changes in the dynamic relation between money growth and the rate of unemployment over time (see Pesaran (1988, p.506)).

3. Data and Sample Periods

Equations $\{(3), (4), (5)\}$ and $\{(6), (4), (5)\}$ comprise the three—equation New Classical system. In this paper, the three equations incorporating the cross—equation restrictions are estimated by maximum likelihood for the periods 1946—73 and 1946—85. It has become common practice in the literature dealing with unobserved variables to use 2SE and M2SE rather than maximum likelihood to estimate the parameters of the system of

equations. In this context, when equations (4) and (5) are first estimated to derive OLS residuals for use in equations (3) or (6), the M2SE of the coefficients of (3) or (6) will not be efficient and typically will not yield consistent estimators of the standard errors.

When M2SE is used, equations (3) and (6) are estimated over 1946–73 and 1946–85, equation (4) is estimated over 1941–73 and 1941–85, and equation (5) is estimated over 1943–73 and 1943–85 (see Barro (1977), Pesaran (1982, 1988) and Rush and Waldo (1988) for details). The reason for the choice of sample periods is not immediately obvious from reading the papers. Barro (1977) estimated an unemployment equation for 1946–73 and a money growth rate equation for 1941–73. Rush and Waldo (1988) and Pesaran (1982, 1988) also use these time periods. Moreover, these latter authors do not re–estimate the rate of money growth equation to adjust for expectations of real federal government expenditure relative to its normal level; Pesaran (1982, p.547) makes an adjustment to the residuals of the Barro (1977) rate of money growth equation to take account of this requirement. Pesaran (1982) also estimates the rate of money growth equation over the period 1942–73, while Rush and Waldo (1988, p.500, footnote 2) use data for 1943–73.

4. Variable Addition Tests

When unobserved variables in New Classical models are replaced by generated regressors, the resulting errors become heteroskedastic and serially correlated. For this reason, non–nested tests based on the assumption of spherical errors will generally be biased for testing the New Classical model as the null against the Keynesian alternative. Moreover, variable addition diagnostic tests based on M2SE may yield invalid inferences because the standard errors will not be estimated consistently.

Pagan (1984, Theorem 8) showed that the estimated standard errors in models estimated by 2SE are no greater than the true standard errors, so that test statistics based on 2SE are generally biased towards rejecting the relevant null hypothesis (see also Murphy and Topel (1985)). An extension of this result to M2SE is given in Appendix A. Since two

of the diagnostic tests used at the single–equation level, namely the RESET test for functional form misspecification of Ramsey (1969, 1974) and the test for serial correlation due to Godfrey (1978) and Breusch and Godfrey (1981), generally exhibit this bias, they need to be recalculated when the relevant null hypothesis is rejected. It is straightforward to show that the variable addition test for serial correlation based on M2SE is not biased when the equation generating the expectations contains only exogenous regressors. However, since virtually all examples available in the literature, including the DM and DG equations in (4) and (5), have lagged values of the dependent variable in the set of regressors, this exception is of little practical interest.

Variable addition non–nested tests of the New Classical model are also biased towards rejection of the null. Since the New Classical model is rejected quite often on the basis of non–nested tests (see Pesaran (1982, 1988)), the combination of the bias of the tests and the empirical evidence towards rejection would seem to reinforce the need to recalculate the test statistics correctly. The mean– and variance–adjusted Cox and Wald–type tests of Godfrey and Pesaran (1983), which are small sample refinements of the Cox test of Pesaran (1974), are asymptotically equivalent under the null hypothesis and under local alternatives to two variable addition non–nested tests, namely the J test of Davidson and MacKinnon (1981) and the JA test of Fisher and McAleer (1981). It is not presently known if this asymptotic equivalence holds in all cases involving models with generated regressors but, if it does, the direction of bias is the same. In such models, the variable addition J and JA tests are biased towards rejection of the null using M2SE since the test statistics are calculated on the basis of an understated covariance matrix. However, since the adjusted Cox and Wald–type tests are based on the ratios of sums of estimated error variances, it is not clear whether these tests are biased and, if so, in which direction. What can be stated is that the original Cox test, being based on the mean–corrected difference of the log–likelihood values of the two models, is not correctly computed for the New Classical null model because it does not take account of the inherent

heteroskedasticity and serial correlation of the errors.

Although single—equation variable addition non—nested tests of the Keynesian model are valid, higher power might be expected by using the New Classical model with cross—equation restrictions imposed as the alternative. In addition, strict comparability with the tests of the New Classical model will be maintained by using the same comprehensive system test procedure within a systems context. However, given the structure of the models, two variable addition non—nested tests of the Keynesian model as the null do not require maximum likelihood estimation of the system at the final stage.

5. Empirical Results

5.1 Estimation

This section presents the results of empirical estimation of the New Classical models as well as the non—nested test statistics of the New Classical and Keynesian models. The maximum likelihood estimates of the original and revised New Classical models are given in Tables 1 and 2, the diagnostic tests for each of the three equations comprising the New Classical system are presented in Table 3, the appropriate diagnostic tests of the New Classical system and tests of various parametric restrictions are given in Table 4, and the results of non—nested tests of the New Classical and Keynesian models against each other using M2SE and maximum likelihood methods are displayed in Tables 5 and 6, respectively.

Since the unemployment equation of the New Classical system is to be compared directly with its Keynesian counterpart, the relevant OLS estimates of the original and revised Keynesian unemployment equations are given in equation (1) and Appendix Table 1 (pages 505 and 507, respectively) of Pesaran (1988). It is worth emphasizing the conformity of signs and magnitudes with prior expectations as well as the statistical significance of most of the estimated coefficients in both versions of the Keynesian specification, and the satisfactory diagnostic test statistics. However, as in Pesaran (1982),

the estimated coefficients of the minimum wage variable are consistently negative, but it is barely significant in the original version in Pesaran (1988). Moreover, the minimum wage variable is deleted in the revised Keynesian model for 1946–85 in Pesaran (1988) since it is not statistically significant.

For purposes of direct comparison with the maximum likelihood estimates presented here, it is helpful to summarize the existing 2SE and M2SE results. Since Barro (1977, 1979) and Small (1979) maintain the assumption that the $FEDV_t$ variable can be anticipated perfectly at time t–1, they do not have an equation for the growth of real federal government expenditure. Hence, their equation for money growth is not estimated efficiently by OLS even if their assumption is warranted and the disturbances of the money growth and unemployment equations are uncorrelated. The unemployment equation is not efficiently estimated by 2SE and the standard errors are not correct. When the unrealistic assumption regarding $FEDV_t$ is relaxed, as in Pesaran (1982, 1988) and Rush and Waldo (1988), the government expenditure growth equation is not estimated efficiently by OLS relative to estimation of the system by maximum likelihood even if the disturbances of the three equations are uncorrelated. The money growth and unemployment equations are not efficiently estimated by M2SE and the calculated standard errors are not correct (see Appendix A for further details).

The government expenditure growth equation of Pesaran (1982) and Rush and Waldo (1988) have all estimated coefficients of the expected signs and are statistically significant; in particular, the lagged unemployment rate has a positive and significant estimated coefficient. Barro's (1977, p. 104) money growth equation has all its estimated coefficients being positive, but the coefficient of lagged growth is not significant. The equivalent equation with $FEDV_t$ replaced by $E_{t-1}(FEDV_t)$ is not given in Pesaran (1982, 1988) or Rush and Waldo (1988), but the estimates (not reported here) for the period 1943–73 are not qualitatively different from those using $FEDV_t$ for 1941–73. Finally, the unemployment equation seems to be quite adequate as far as determination of signs and

magnitudes is concerned and, with the qualification that the standard errors are understated, most coefficients seem to be "statistically significant". The consistent exception to the general result is the estimated coefficient of the minimum wage variable, which seems to be highly sensitive both in sign and magnitude to the specification used. However, since the estimated coefficients typically have t-ratios that are below conventional levels in spite of their being biased upwards, there would seem to be little of real concern about this variable.

The coefficients in Tables 1 and 2 generally have the same signs and similar orders of magnitude as their M2SE counterparts, the exception being the lagged unemployment variable in the government expenditure growth equation, where the maximum likelihood estimate is consistently negative but insignificant. For both sample periods, the minimum wage variable has positive but insignificant estimated coefficients for the original New Classical model and negative but insignificant coefficients for the revised model. The time trend and the lagged fiscal shock are less significant than they might appear on the basis of M2SE for the period 1946–73 (see Pesaran (1982, Table 5)), but the time trend is statistically significant in the revised New Classical model estimated by maximum likelihood for 1946–85.

It is worth mentioning that, while the estimated standard errors obtained by M2SE on computer packages are understated relative to the correct (but inefficient) M2SE standard errors using the formula in Theorem 4 of Appendix A, maximum likelihood is (asymptotically) more efficient than M2SE and, hence, should yield smaller standard errors in large samples than the correct M2SE standard errors. Although not reported here, the correct M2SE standard errors are generally much larger than their maximum likelihood counterparts. However, it is not obvious whether the maximum likelihood estimates should have smaller estimated standard errors than their (understated) M2SE counterparts based on the incorrect formula (as are presented in all of the papers mentioned above). For example, Murphy and Topel (1985, Table 1, p.372) report the understated 2SE, the correct

(but inefficient) 2SE and maximum likelihood estimates of the parameters of Barro's (1977) original unemployment equation as part of the basic two–equation system, together with the corresponding standard errors, using data for 1946–73. The maximum likelihood standard errors are always smaller than the correct 2SE standard errors, sometimes substantially, and are even less than the understated 2SE standard errors for two of the six estimated coefficients.

5.2 Diagnostic and Hypothesis Tests

The results of four diagnostic tests for each equation of both versions of the New Classical system are provided for both sample periods in Table 3. Descriptions of each test and the methods of calculation in a systems context are described in Appendix B. On the basis of recent Monte Carlo evidence for linear regression models in Godfrey et al. (1988) and Thursby (1989), the most powerful version of the RESET test was adopted by using the squared fitted values of each dependent variable. The serial correlation test should be powerful against any alternative hypothesis exhibiting at least first–order autoregressive or moving average characteristics because annual data are used (Pesaran (1988, p.505) also tested against a first–order alternative). The tests for heteroskedasticity and normality are based on the Lagrange multiplier principle. In the calculation of each of these tests, it is presumed that only the equation being tested might be departing from the assumed conditions of the null hypothesis.

Apart from a significant value of RESET at the five percent level for the money growth equation in the revised model for 1946–73, no significant functional form misspecification, serial correlation or heteroskedasticity is detected in any of the three equations comprising the original or revised New Classical systems for either sample period. However, the government expenditure growth equation exhibits substantial non–normal errors. Since the marginal distribution of the errors in one of the three equations is not normally distributed, the errors of the New Classical system cannot be

jointly normally distributed. However, the use of the rational expectations hypothesis is not conditional on joint normality of the errors, so the observed non—normality should not be viewed as an empirical rejection of the New Classical model. Moreover, these diagnostic test results are in general agreement with those given in Pesaran (1988).

It is worth reiterating that the M2SE method used by Pesaran (1988) involves serially correlated and heteroskedastic errors in both the money growth and unemployment equations. Since the diagnostic tests generally used for serial correlation and heteroskedasticity are not designed specifically for the types of error structures inherent in models using M2SE methods, it is possible that non—detection of certain problems by M2SE reflects low power of the tests used rather than an absence of the problems being investigated. Moreover, although tests of heteroskedasticity and other tests based on even moments are not affected by the presence of generated regressors because the use of squared residuals eliminates any parameter estimation effects, this is not the case for tests based on odd moments. Thus, the joint test of normality based on the third and fourth moments is affected by generated regressors (see Pagan and Hall (1983a) for further details).

Diagnostic tests for functional form misspecification and serial correlation for the New Classical system are presented in Table 4, and there appears to be no evidence of significant departures from the null hypothesis in either case. Tests of three sets of parametric restrictions are also given in Table 4. The cross—equation restrictions (see Mishkin (1983, Section 2.2) and Pesaran (1987, Section 7.5)) are also supported by the data, but it should be stressed that, given the low degrees of freedom involved, the powers of such tests are likely to be quite low for the problem considered here, especially for the 1946—73 sample period. When the anticipated components are added to the appropriate New Classical model, they are found not to be statistically significant. In answer to the question posed by Mishkin (1982), namely "Does anticipated monetary policy matter?", the answer using Barro's (1977) original annual data and an updated annual version is

resoundingly in the negative, although Mishkin answered in the affirmative using seasonally adjusted, U.S. quarterly data for 1954—76. Finally, the unanticipated components are highly significant in both versions of the New Classical model for both sample periods, so that monetary shocks do seem to matter in explaining U.S. unemployment.

Using the data set for 1946—73 and Barro's (1977) original two—equation New Classical system based on the assumption that FEDV can be anticipated perfectly, Liederman (1980) uses maximum likelihood estimation to examine if unanticipated money growth affects unemployment. It is found that the rational expectations (or overidentifying) restrictions, the restrictions implied by the 'structural neutrality' hypothesis, and the restrictions implied by the joint hypothesis of the two just mentioned are all supported by the data. Thus, it would seem that money growth affects U.S. unemployment only through its unanticipated, and not its anticipated, component.

5.3 Non—nested Tests

In an early attempt to choose between competing non—nested models as well as to test them against each other, Barro (1977, pp. 108—109) examined two non—nested alternatives to his own New Classical specification. Three alternative definitions of the money stock were used to generate three alternative series of money supply shocks and then, conditional upon the New Classical framework, the model yielding the highest coefficient of determination in explaining unemployment was chosen as the best. A far more interesting development arose when he tested the anticipated and unanticipated components of monetary policy against each other by testing exclusion restrictions within a more general model. Taking the anticipated and unanticipated versions as two non—nested alternatives, Barro's procedure may be interpreted as testing a null hypothesis by comparing two estimators of selected parameters of interest of the non—nested alternative model. In this context, Deaton (1982), Dastoor (1983) and Gourieroux, Monfort and

Trognon (1983) derived a non—nested F test based on selected parameters of interest, and this may be made operational by using the pseudo—true values of the selected parameters. McAleer and Pesaran (1986) showed that a similar analysis could be conducted using Roy's union—intersection principle, while Mizon and Richard (1986) derived an identical F test to those mentioned previously based on the encompassing principle.

Barro (1977, p.109) found that the anticipated component of monetary policy was not statistically significant whereas the unanticipated component was statistically significant. However, as shown in Pagan (1984), the tests conducted by Barro are biased towards rejection of the null hypothesis in each case because the estimated standard errors are biased downwards. Thus, while Barro's result concerning the insignificance of the anticipated component cannot be overturned by a correctly computed test statistic, the same might not be true for the unanticipated component.

The same reservations might need to be directed at the empirical evidence reported in Pesaran (1988) regarding the superiority of the Keynesian model of unemployment relative to Rush and Waldo's (1988) extension of Barro's (1977) New Classical model. Table 5 presents the results of five non—nested tests based on M2SE. The variable addition J, JA and F tests obtained as standard output on computer packages are biased towards rejection of the New Classical model when it is the null and, if the adjusted Cox test or the Wald—type test, $\bar{N}$ and W, respectively, are asymptotically equivalent to these tests, the direction of bias is the same. Test statistics for the Keynesian null are valid in all cases since each of the explanatory variables is directly measurable. On the basis of the calculated statistics, it is clear why the Keynesian model might be seen to be superior to its New Classical counterpart. Whenever the Keynesian model is the null it is not rejected by its New Classical competitor. Only when the revised New Classical model is the null for the 1946—85 sample period can it be safely determined that the null is not rejected against the Keynesian alternative, since the decision cannot be overturned by a correct calculation of the test statistics. In other cases of rejection of the New Classical model, judgment

needs to be suspended in view of the upward bias of the variable addition non–nested tests. Moreover, the J test is known to have a penchant for over–rejecting a true null hypothesis in small samples relative to the predictions of asymptotic theory (even when the standard errors are not biased downwards), while the JA and F tests are known to have lower power than the other available tests (for further details, see Davidson and MacKinnon (1982), Godfrey and Pesaran (1983), and King and McAleer (1987)).

Table 5 also presents, in square brackets, the correct variable addition non–nested J, JA and asymptotic F test statistics for the New Classical models using the formula in Theorem 4 of Appendix A. In all cases, the correctly calculated test statistics using (inefficient) M2SE are smaller, sometimes substantially, than their counterparts obtained using the understated standard errors. What is of particular interest in light of the debate between Pesaran (1982, 1988) and Rush and Waldo (1988) is that none of the New Classical models is rejected against the Keynesian alternative at conventional levels of significance using the correct formula.

Since the previous rejections of the New Classical model in the literature based on M2SE using the incorrect standard errors would appear to be suspect, the variable addition non–nested J, JA and asymptotic F tests based on maximum likelihood estimation are reported in Table 6. The Keynesian null hypothesis is not rejected against the New Classical alternative, thereby adding further support to Pesaran's results on the validity of the Keynesian specification. However, when the New Classical model is the null, the outcome depends on the test used and, in one case, also on the level of significance used. The J and JA tests are in agreement concerning rejection of the New Classical null in three of the four cases, with the asymptotic F test indicating non–rejection in all cases. Only in the case of the revised New Classical model as the null do the JA and asymptotic F tests agree with each other, with the J test indicating rejection at the five percent level. Therefore, the variable addition non–nested test statistics calculated by maximum likelihood lend support to Pesaran's (1988) result concerning rejection of the New Classical

model but not the Keynesian model if the J and JA tests are used rather than the asymptotic F test. However, an improved version of the New Classical model can withstand the challenge of the Keynesian model, even though it cannot itself reject the Keynesian explanation of unemployment in the U.S.

6. Conclusion

In this paper several Keynesian and New Classical models of unemployment for the U.S. are re–evaluated. Since two step estimation (2SE) and multivariate two step estimation (M2SE) are generally neither efficient nor provide consistent estimators of the standard errors for the New Classical models of unemployment available in the literature, maximum likelihood methods are used for estimating and testing the New Classical models. The adequacy of both the Keynesian and New Classical models is tested by the use of diagnostic and non–nested tests, and several parametric restrictions are also tested for the three–equation New Classical system. Although the existing empirical results in the literature using 2SE and M2SE would seem to favour strongly the Keynesian specification over the New Classical system, two important findings of this paper are that neither specification is rejected on the basis of correctly calculated (though inefficient) variable addition non–nested test statistics, and that an improved version of the New Classical system is not rejected against the Keynesian alternative when estimation and testing are undertaken within a systems context.

TABLE 1

Maximum Likelihood Estimates of New Classical Models, 1946–73

Dependent Variable	Explanatory Variable	Original Model			Revised Model		
		Coefficient Estimate	Standard Error	t–ratio	Coefficient Estimate	Standard Error	t–ratio
DG_t	Intercept	−0.058	0.161	−0.360	−0.035	0.157	−0.223
	DG_{t-1}	0.301	0.059	5.102	0.293	0.059	4.966
	UN_{t-1}	−0.035	0.052	−0.673	−0.028	0.051	−0.549
	WAR_t	−0.142	0.011	−12.909	−0.139	0.013	−10.692
DM_t	Intercept	0.093	0.021	4.429	0.081	0.021	3.857
	DM_{t-1}	0.463	0.119	3.891	0.406	0.128	3.172
	DM_{t-2}	0.123	0.101	1.218	0.163	0.108	1.509
	UN_{t-1}	0.028	0.007	4.000	0.024	0.007	3.429
	$E_{t-1}(FEDV_t)$	0.066	0.011	6.000	0.069	0.013	5.308
UN_t	Intercept	−2.839	0.197	−14.411	−2.854	0.173	−16.497
	MIL_t	−4.788	0.957	−5.003	−4.148	1.025	−4.047
	$MINW_t$	0.200	0.534	0.375	−0.587	0.796	−0.737
	$DMRH_t$	−4.056	1.941	−2.090	−3.843	1.899	−2.024
	$DMRH_{t-1}$	−11.750	1.844	−6.372	−11.662	1.790	−6.515
	$DMRH_{t-2}$	−5.612	2.228	−2.519	−5.998	2.382	−2.518
	t				0.010	0.007	1.429
	DGR_{t-1}				0.478	0.411	1.163

Note: The t–ratios have been rounded to correspond to the coefficient estimates and their standard errors being reported to three decimal places.

TABLE 2

Maximum Likelihood Estimates of New Classical Models, 1946–85

Dependent Variable	Explanatory Variable	Original Model			Revised Model		
		Coefficient Estimate	Standard Error	t–ratio	Coefficient Estimate	Standard Error	t–ratio
DG_t	Intercept	−0.060	0.085	−0.706	−0.051	0.089	−0.573
	DG_{t-1}	0.307	0.051	6.020	0.300	0.052	5.769
	UN_{t-1}	−0.036	0.029	−1.241	−0.032	0.030	−1.067
	WAR_t	−0.140	0.009	−15.556	−0.139	0.010	−13.900
DM_t	Intercept	0.108	0.012	9.000	0.092	0.013	7.077
	DM_{t-1}	0.391	0.106	3.689	0.328	0.116	2.828
	DM_{t-2}	0.221	0.090	2.456	0.267	0.095	2.811
	UN_{t-1}	0.034	0.004	8.500	0.029	0.005	5.800
	$E_{t-1}(FEDV_t)$	0.070	0.011	6.364	0.071	0.012	5.917
UN_t	Intercept	−2.904	0.193	−15.047	−2.976	0.177	−16.814
	MIL_t	−5.129	0.969	−5.293	−3.812	0.990	−3.851
	$MINW_t$	0.641	0.462	1.387	−0.638	0.597	−1.069
	$DMRH_t$	−5.023	1.755	−2.862	−4.248	1.489	−2.853
	$DMRH_{t-1}$	−11.029	1.725	−6.394	−10.692	1.519	−7.039
	$DMRH_{t-2}$	−5.458	2.071	−2.635	−5.934	1.980	−2.997
	t				0.016	0.006	2.667
	DGR_{t-1}				0.506	0.366	1.383

Note: The t–ratios have been rounded to correspond to the coefficient estimates and their standard errors being reported to three decimal places.

TABLE 3

Diagnostic Tests of the Equations Comprising the New Classical Models
Calculated by Maximum Likelihood

Sample Period	Equation	Model	RESET	Serial Correlation	Heteroskedasticity	Normality
				Diagnostic Tests		
1946–73	DG	Original	0.43(1)	0.57(1)	0.19(1)	25.00**(2)
		Revised	0.12(1)	0.56(1)	0.21(1)	24.67**(2)
	DM	Original	2.45(1)	1.97(1)	0.003(1)	1.31(2)
		Revised	3.48(1)	2.84(1)	0.02(1)	0.66(2)
	UN	Original	3.82(1)	0.63(1)	0.18(1)	1.08(2)
		Revised	3.98*(1)	0.19(1)	0.003(1)	1.31(2)
1946–85	DG	Original	0.04(1)	0.86(1)	0.14(1)	75.20**(2)
		Revised	0.35(1)	0.83(1)	0.14(1)	75.33**(2)
	DM	Original	1.07(1)	0.54(1)	0.81(1)	0.36(2)
		Revised	0.06(1)	0.73(1)	0.04(1)	0.01(2)
	UN	Original	0.54(1)	1.38(1)	0.68(1)	5.67(2)
		Revised	1.18(1)	1.30(1)	2.08(1)	1.66(2)

Notes: 1 Degrees of freedom for the asymptotic chi–squared tests are given in parentheses immediately following the calculated statistic. The RESET and serial correlation tests are likelihood ratio tests, while the heteroskedasticity and normality tests are Lagrange multiplier tests.

    * Denotes statistically significant at the five percent level.

    ** Denotes statistically significant at the one percent level.

TABLE 4

Tests of the New Classical Systems Calculated by Maximum Likelihood

Sample Period	Model	Tests of Parametric Restrictions			Diagnostic Tests	
		Cross–equation Restrictions	Anticipated Components	Unanticipated Components	RESET	Serial Correlation
1946–73	Original	21.75(18)	4.96(3)	50.03*(3)	5.11(3)	3.32(3)
	Revised	20.69(17)	6.19(4)	48.97*(4)	3.28(3)	3.52(3)
1946–85	Original	22.72(18)	6.68(3)	58.17*(3)	3.33(3)	2.64(3)
	Revised	19.07(17)	6.06(4)	64.10*(4)	0.15(3)	2.80(3)

Notes: 1 Degrees of freedom for the asymptotic chi–squared tests are given in parentheses immediately following the calculated statistic. All tests are likelihood ratio tests.

      * Denotes statistically significant at the five percent level.

TABLE 5

Non—nested Tests Based on Multivariate Two Step Estimation

Null Model	Alternative Model	Sample Period	Non—nested Tests				
			N	W	J	JA	F
Original New Classical	Original Keynesian	1946—73	−3.33	−2.42	4.49 [1.44]	2.62 [1.98]	3.42(5,17) [1.77(5,17)]
Original Keynesian	Original New Classical	1946—73	−0.03	−0.03	0.60	−0.19	0.98(3,17)
Revised New Classical	Original Keynesian	1946—73	−2.45	−1.93	3.09 [1.04]	2.40 [1.45]	2.14(4,16) [1.12(4,16)]
Original Keynesian	Revised New Classical	1946—73	−0.17	−0.17	0.93	0.05	0.72(4,16)
Original New Classical	Revised Keynesian	1946—85	−3.88	−2.98	4.02 [1.55]	3.55 [2.15]	2.74(6,28) [1.36(6,28)]
Revised Keynesian	Original New Classical	1946—85	−0.38	−0.37	0.54	0.45	0.59(4,28)
Revised New Classical	Revised Keynesian	1946—85	−1.25	−1.15	1.88 [0.72]	1.36 [0.68]	0.75(5,27) [0.55(5,27)]
Revised Keynesian	Revised New Classical	1946—85	−1.02	−0.96	1.62	1.28	0.58(5,27)

Notes: 1. The degrees of freedom for the F test statistics are given in parentheses immediately following the calculated statistics. All other tests are asymptotically distributed under the null hypothesis as $N(0,1)$. The non—nested test statistics were computed using the computer package Microfit (see Pesaran and Pesaran (1989)).

2. When the New Classical model is the null, the variable addition J, JA and F test statistics based on M2SE are biased towards rejection of the null hypothesis. If the N and W tests are asymptotically equivalent to the J and JA test statistics under the null and under local alternatives, the direction of bias of the N and W tests is the same.

3. The calculated test statistics given in square brackets are based on the correct M2SE covariance matrix (see Theorem 4 of Appendix A).

TABLE 6

Variable Addition Non—nested Tests Calculated by Maximum Likelihood

Null Model	Alternative Model	Sample Period	Non—nested Tests		
			J	JA	Asymptotic F
Original New Classical	Original Keynesian	1946–73	8.78**(1)	8.94**(1)	11.04(5)
Original Keynesian	Original New Classical	1946–73	1.04	0.19	3.60(3)
Revised New Classical	Original Keynesian	1946–73	8.15**(1)	6.03**(1)	8.34(4)
Original Keynesian	Revised New Classical	1946–73	1.45	0.26	3.74(4)
Original New Classical	Revised Keynesian	1946–85	9.72**(1)	9.10**(1)	11.94(6)
Revised Keynesian	Original New Classical	1946–85	0.49	0.37	3.60(4)
Revised New Classical	Revised Keynesian	1946–85	4.10*(1)	2.08(1)	5.40(5)
Revised Keynesian	Revised New Classical	1946–85	1.44	0.77	4.04(5)

Notes: 1 Degrees of freedom for the chi–squared test statistics are given in parentheses. When the Keynesian model is the null, the J and JA test statistics are asymptotically distributed as $N(0, 1)$.

* Denotes statistically significant at the five percent level.

** D^notes statistically significant at the one percent level.

Appendix A

Multivariate Two Step Estimation of the Revised New Classical Model

Using the notation of Pagan (1984) and McAleer and McKenzie (1988), the Revised New Classical model given in equations (6), (4) and (5) can be written in matrix form, respectively, as

$$y = \eta\gamma_1 + \eta_{-1}\gamma_2 + \eta_{-2}\gamma_3 + v_{-1}\pi + X\beta + e \tag{A1}$$

$$z_1 = W_1\alpha_1 + (FEDV - 0.8v)\alpha_2 + \eta \tag{A2}$$

$$z_2 = W_2\psi + v \tag{A3}$$

in which $y = UN$, $\eta_{-i} = DMRH_{-i}$ ($i = 0, 1, 2$) and $\eta_0 = \eta$, $v_{-1} = DGR_{-1}$, $X = [1 : MIL : MINW : t]$, $z_1 = DM$, $W_1 = [1 : DM_{-1} : DM_{-2} : UN_{-1}]$, $v = DGR$, $z_2 = DG$, $W_2 = [1 : DG_{-1} : UN_{-1} : WAR]$, and the errors e, η and v are independently and identically distributed random variables with zero means and variances σ_e^2, σ_η^2 and σ_v^2, respectively.

Equations (A2) and (A3) comprise a two–equation expectations system which may be estimated by OLS/2SE or maximum likelihood. For purposes of estimation, equation (A2) may be rewritten as

$$z_1 = W_1\alpha_1 + (FEDV - 0.8\hat{v})\alpha_2 + \eta + (v - \hat{v})\alpha_2^* = \Phi\alpha + u \tag{A4}$$

in which $\Phi = [W_1 : (FEDV - 0.8\hat{v})]$, $\alpha_2^* = -0.8\alpha_2$, $\alpha = (\alpha_1, \alpha_2)'$, $\hat{v} = M_2z_2 = M_2v$, $v - \hat{v} = (I - M_2)v$, $M_2 = I - W_2(W_2'W_2)^{-1}W_2'$ and $u = \eta + (v - \hat{v})\alpha_2^* = \eta + \alpha_2^*(I - M_2)v$. The 2SE results on efficiency and consistent estimation of standard errors are available in Pagan (1984). To summarize, 2SE of equation (A4) is not efficient unless W_1 and W_2 are orthogonal or W_1 appears in W_2, by an application of Theorems 4 and 7(i) in Pagan (1984) (for a very simple alternative proof, see McAleer and McKenzie (1988)).

However, given the definitions of W_1 and W_2, neither of these conditions is satisfied here so 2SE is not efficient. The error variance σ_η^2 is estimated consistently by 2SE, as is shown for completeness in Theorem 1 below, although the result is implied in Pagan (1984) and assumed in Murphy and Topel (1985). Finally, the 2SE standard errors are generally understated (see Theorem 8 in Pagan (1984) and Theorem 1 in Murphy and Topel (1985)). It also follows that diagnostic and non–nested tests based on variable addition and 2SE are generally biased towards rejection of the null hypotheses.

THEOREM 1. *The estimated error variance from equation (A4) using OLS/2SE is a consistent estimator of σ_η^2.*

PROOF. From equation (A4), $z_1 = \Phi\alpha + u$ so that the OLS estimator of the error variance is

$$T^{-1}\hat{u}'\hat{u} = T^{-1}u'u - T^{-1}u'\Phi(\Phi'\Phi)^{-1}\Phi'u$$

where $\Phi'u = \begin{bmatrix} W_1'u \\ (FEDV - 0.8\hat{v})'u \end{bmatrix}$.

Given $\hat{v} = M_2 v$ and $u = \eta + \alpha_2^*(I - M_2)v$, it follows that $T^{-1}W_1'u \xrightarrow{P} 0$,

$T^{-1}FEDV'\eta \xrightarrow{P} 0$, $T^{-1}FEDV'v \xrightarrow{P} 0$ and $T^{-1}\hat{v}'u \xrightarrow{P} 0$, so that $T^{-1}\Phi'u \xrightarrow{P} 0$

and $(T^{-1}\hat{u}'\hat{u} - T^{-1}u'u) \xrightarrow{P} 0$. Since

$$T^{-1}u'u = T^{-1}\eta'\eta + 2\alpha_2^*T^{-1}\eta'(I - M_2)v + T^{-1}\alpha_2^{*2}v'(I - M_2)v,$$

$T^{-1}\eta'\eta \xrightarrow{P} \sigma_\eta^2$, $T^{-1}\eta'(I - M_2)v \xrightarrow{P} 0$ and $T^{-1}v'(I - M_2)v \xrightarrow{P} 0$, it follows that

$T^{-1}u'u \xrightarrow{P} \sigma_\eta^2$ and $T^{-1}\hat{u}'\hat{u} \xrightarrow{P} \sigma_\eta^2$. ∎

Equations (A1) — (A3) comprise a three—equation system, namely a univariate structural equation with a two—equation expectations system. For purposes of estimation, equation (A1) may be rewritten as

$$y = \hat{\eta}\gamma_1 + \hat{\eta}_{-1}\gamma_2 + \hat{\eta}_{-2}\gamma_3 + \hat{v}_{-1}\pi + X\beta + e + (\eta - \hat{\eta})\gamma_1$$

$$+ (\eta_{-1} - \hat{\eta}_{-1})\gamma_2 + (\eta_{-2} - \hat{\eta}_{-2})\gamma_3 + (v_{-1} - \hat{v}_{-1})\pi$$

or

$$y = Q\Theta + \xi \qquad (A5)$$

in which $Q = [\hat{\eta} : \hat{\eta}_{-1} : \hat{\eta}_{-2} : \hat{v}_{-1} : X]$, $\Theta = (\gamma_1, \gamma_2, \gamma_3, \pi, \beta')'$ and

$$\xi = e + (\eta - \hat{\eta})\gamma_1 + (\eta_{-1} - \hat{\eta}_{-1})\gamma_2 + (\eta_{-2} - \hat{\eta}_{-2})\gamma_3 + (v_{-1} - \hat{v}_{-1})\pi. \qquad (A6)$$

It is necessary to derive $E(\xi\xi')$ to enable inferences to be drawn from M2SE of equation (A5). Defining $\Phi_{-i} = [W_{1, -i} : FEDV_{-i} - 0.8\hat{v}_{-i}]$ and $\hat{z}_{1, -i} = \Phi_{-i}(\Phi'\Phi)^{-1}\Phi'z_1$ for $i = 0, 1, 2$, it follows that

$$\hat{\eta}_{-i} = z_{1, -i} - \hat{z}_{1, -i} = u_{-i} - \Phi_{-i}(\Phi'\Phi)^{-1}\Phi'u$$

or

$$\hat{\eta}_{-i} = \eta_{-i} + (v_{-i} - \hat{v}_{-i})\alpha_2^* - \Phi_{-i}(\Phi'\Phi)^{-1}\Phi'u. \qquad (A7)$$

Since

$$v_{-i} - \hat{v}_{-i} = W_{2, -i}(W_2'W_2)^{-1}W_2'v \qquad (A8)$$

substitution of (A8) into (A7) yields

$$\eta_{-i} - \hat{\eta}_{-i} = -W_{2, -i}(W_2'W_2)^{-1}W_2'v\,\alpha_2^* + \Phi_{-i}(\Phi'\Phi)^{-1}\Phi'(\eta + \alpha_2^*(I-M_2)v)$$

or

$$\eta_{-i} - \hat{\eta}_{-i} = \Phi_{-i}(\Phi'\Phi)^{-1}\Phi'\eta + \alpha_2^*[\Phi_{-i}(\Phi'\Phi)^{-1}\Phi'(I-M_2) - W_{2, -i}(W_2'W_2)^{-1}W_2']v. \quad (A9)$$

Substitution of (A8) and (A9) into (A6) enables ξ to be rewritten as

$$\xi = e + S_1\eta + \alpha_2^* S_2 v \tag{A10}$$

in which

$$S_1 = (\gamma_1\Phi + \gamma_2\Phi_{-1} + \gamma_3\Phi_{-2})(\Phi'\Phi)^{-1}\Phi' \tag{A11}$$

$$S_2 = S_1(I-M_2) - [\gamma_1 W_2 + (\gamma_2 - \pi/\alpha_2^*)W_{2,-1} + \gamma_3 W_{2,-2}](W_2'W_2)^{-1}W_2'. \tag{A12}$$

The covariance matrix of ξ, which is required for analysing the efficiency of M2SE and the bias in the covariance matrix of the M2SE of Θ in (A5), is given in the following lemma.

LEMMA 1. $E(\xi\xi') = V = \sigma_e^2 I + \sigma_\eta^2 S_1 S_1' + \alpha_2^{*2}\sigma_v^2 S_2 S_2'$.

PROOF. Since e, η and v are independent, by assumption, the covariance matrix of ξ is the sum of the covariance matrices of each of the three terms on the right–hand side of (A10). ∎

Although several alternative equivalent forms of the necessary and sufficient condition for efficiency of least squares estimators among single equation estimators have been developed independently by several authors (see McAleer (1989) for further details), the method of proof used here extends the analysis of McAleer and McKenzie (1988) for 2SE based on the results of Kruskal (1968). The appropriate condition in terms of M2SE of the parameters of (A5) is summarized in the following theorem.

THEOREM 2. *The M2SE of Θ in equation (A5) is efficient if and only if there exists a matrix F such that*

$$VQ = QF$$

where V is defined in Lemma 1. ∎

The result regarding the efficiency of M2SE is given in the following theorem.

THEOREM 3. *The M2SE of Θ in equation ($A5$) is inefficient unless Q is contained in or is orthogonal to each of Φ, Φ_{-1}, Φ_{-2}, W_2, $W_{2,-1}$ and $W_{2,-2}$.*

PROOF. Substitution of (A11) and (A12) into the expression for V in Lemma 1 shows that the necessary and sufficient condition of Theorem 2 is not satisfied unless $S_1 S_1' Q$ and $S_2 S_2' Q$ are either linear combinations of Q or are null matrices. Thus, M2SE is inefficient unless Q is contained in or is orthogonal to each of Φ, Φ_{-1}, Φ_{-2}, W_2, $W_{2,-1}$ and $W_{2,-2}$. ∎

However, since neither of the exceptions given in Theorem 3 holds for the problem considered here, M2SE is not efficient.

Denoting the true covariance matrix of the M2SE of Θ in equation (A5) as $(Q'Q)^{-1}Q'VQ(Q'Q)^{-1}$, we have the following theorem.

THEOREM 4. *The standard errors computed by applying M2SE to equation ($A5$) are no greater than the true standard errors.*

PROOF. Substitution of V from Lemma 1 into the formula for the true standard errors yields

$$(Q'Q)^{-1}Q'VQ(Q'Q)^{-1} = \sigma_e^2(Q'Q)^{-1} + \sigma_\eta^2(Q'Q)^{-1}Q'S_1S_1'Q(Q'Q)^{-1}$$
$$+ \alpha_2^{*2}\sigma_v^2(Q'Q)^{-1}Q'S_2S_2'Q(Q'Q)^{-1}$$

which, by virtue of the positive semi–definiteness of the second and third terms, exceeds the computed M2SE standard errors, $\sigma_e^2(Q'Q)^{-1}$. ∎

Although the computed M2SE covariance matrix is given by $\sigma_e^2(Q'Q)^{-1}$, it is necessary to prove that the error variance in (A5) estimated by M2SE is consistent for σ_e^2. Some preliminary results are given in Lemmas $2 - 4$.

LEMMA 2. $T^{-1}\xi'\xi \xrightarrow{P} \sigma_e^2$.

PROOF. Using equation (A10), it follows that

$$T^{-1}\xi'\xi = T^{-1}e'e + T^{-1}\eta'S_1'S_1\eta + \alpha_2^{*2}v'S_2'S_2v + 2e'S_1\eta$$
$$+ 2\alpha_2^*e'S_2v + 2\alpha_2^*\eta'S_1'S_2v.$$

Given the independence of e, η and v, and the results that $T^{-1}W_1'\eta$, $T^{-1}W_2'\eta$, $T^{-1}W_2'v$, $T^{-1}W_{1,-i}'e$, $T^{-1}W_{2,-i}'\eta$ and $T^{-1}W_{2,-i}'e$ (for $i = 0, 1, 2$) all converge in probability to null vectors, then $(T^{-1}\xi'\xi - T^{-1}e'e) \xrightarrow{P} 0$. Since $T^{-1}e'e \xrightarrow{P} \sigma_e^2$, the result follows. ∎

LEMMA 3. (i) $\qquad T^{-1}\Phi_{-i}'v_{-1} \xrightarrow{P} \begin{bmatrix} 0 \\ c_i \end{bmatrix} \quad$ for $i = 0, 1, 2$

$$\text{where } c_i = \begin{cases} -0.8\,\sigma_v^2, \text{ for } i=1 \\ 0, \text{ for } i=0,2 \end{cases}.$$

(ii) $\qquad T^{-1}\Phi_{-i}'\eta_{-j} \xrightarrow{P} \begin{bmatrix} c_{ij} \\ 0 \end{bmatrix} \quad$ for $i, j = 0, 1, 2$

$$\text{where } c_{ij} \begin{cases} \neq 0, \text{ for } i < j \\ = 0, \text{ otherwise} \end{cases}.$$

PROOF. (i) Using the definitions of Φ_{-i} and v_{-1}, it follows that

$$\Phi_{-i}'v_{-1} = \begin{bmatrix} W_{1,-i}'v_{-1} \\ (\text{FED}V_{-i} - 0.8\hat{v}_{-i})'v_{-1} \end{bmatrix}$$

$$= \begin{bmatrix} W_{1,-i}'v_{-1} \\ \text{FEDV}_{-i}'v_{-1} - 0.8\,[v_{-i}'v_{-1} - v'W_2(W_2'W_2)^{-1}W_{2,-i}'v_{-1}] \end{bmatrix}$$

and

$$T^{-1}W_{1,-i}'v_{-1} \xrightarrow{P} 0 \text{ (since } DG_{-1} \text{ does not appear in } W_{1,-i})$$

$$T^{-1}FEDV_{-i}'v_{-1} \xrightarrow{p} 0$$

$$T^{-1}v_{-i}'v_{-1} \xrightarrow{p} \begin{cases} \sigma_v^2, & \text{for } i=1 \\ 0, & \text{for } i=0,2 \end{cases}$$

$$T^{-1}W_2'v \xrightarrow{p} 0$$

$$T^{-1}W_{2,-i}'v_{-1} \xrightarrow{p} \begin{cases} c \neq 0, & \text{for } i=0 \\ 0, & \text{for } i=1,2 \end{cases}$$

(since DG_{-1} appears in W_2 but not in $W_{2,-1}$ or $W_{2,-2}$).

(ii) $\Phi_{-i}'\eta_{-j} = \begin{bmatrix} W_{1,-i}'\eta_{-j} \\ (FEDV_{-i} - 0.8v_{-i})'\eta_{-j} \end{bmatrix}$

$= \begin{bmatrix} W_{1,-i}'\eta_{-j} \\ FEDV_{-i}'\eta_{-j} - 0.8 [v_{-i}'\eta_{-j} - v'W_2(W_2'W_2)^{-1}W_{2,-i}'\eta_{-j}] \end{bmatrix}$

and

$$T^{-1}W_{1,-i}'\eta_{-j} \xrightarrow{p} \begin{cases} c_{ij} \neq 0, & \text{for } i < j \\ 0, & \text{otherwise} \end{cases}$$

(since $W_{1,-i}$ contains DM_{-1-i} and DM_{-2-i})

$$T^{-1}FEDV_{-i}'\eta_{-j} \xrightarrow{p} 0$$

$$T^{-1}v_{-i}'\eta_{-j} \xrightarrow{p} 0 \quad \text{for } i, j = 0, 1, 2$$

$$T^{-1}W_2'v \xrightarrow{p} 0$$

$$T^{-1}W_{2,-i}'\eta_{-j} \xrightarrow{p} 0 \quad \text{for } i, j = 0, 1, 2 \text{ (since } DM_{-j} \text{ does not appear in } W_{2,-i}). \quad \blacksquare$$

LEMMA 4. $T^{-1}Q'\xi \xrightarrow{p} 0.$

PROOF. Given $Q = [\hat{\eta} : \hat{\eta}_{-1} : \hat{\eta}_{-2} : \hat{v}_{-1} : X]$ and $\xi = e + S_1\eta + \alpha_2^* S_2 v$, the result follows from the conditions given in the proof of Lemma 2, the results of Lemma 3 and the assumption $T^{-1}X'e \xrightarrow{p} 0.$ $\quad \blacksquare$

The previous results may now be used to prove the following theorem.

THEOREM 5. *The estimated error variance from equation (A5) using OLS/M2SE is a consistent estimator of σ_e^2.*

PROOF. From equation (A5), $y = Q\Theta + \xi$ so that the OLS estimator of the error variance is

$$T^{-1}\hat{\xi}'\hat{\xi} = T^{-1}\xi'\xi - T^{-1}\xi'Q(Q'Q)^{-1}Q'\xi.$$

The second term on the right–hand side converges to zero in probability by Lemma 4, so that $(T^{-1}\hat{\xi}'\hat{\xi} - T^{-1}\xi'\xi) \xrightarrow{P} 0$. Using Lemma 2, $T^{-1}\hat{\xi}'\hat{\xi} \xrightarrow{P} \sigma_e^2$. ∎

Therefore, the M2SE of the error variance of equation (A5) is consistent for σ_e^2, the true error variance of equation (A1). The results of Theorems 4 and 5 suggest that the standard errors estimated by M2SE are no greater than the true standard errors, so that t–ratios will be biased upwards. It also follows that variable addition diagnostic and non–nested tests are biased towards rejection of the relevant null hypotheses.

There are some exceptions to the general results given in Theorems 3–5. For example, it is possible to show that the M2SE of the coefficient of $\hat{\eta}$, the current unanticipated variable, is efficient (by an extension of Proposition 3.4 in Pagan (1986)) and that its standard error is consistently estimated (by an extension of Proposition 3.3 in Pagan (1986)). However, there would seem to be little practical use in these results since the remaining parameters are inefficient and their estimated standard errors are inconsistent. Moreover, the variable addition diagnostic and non–nested tests are still biased towards rejection of the null hypotheses.

Appendix B

Systems Estimation and Testing of the Original New Classical Model for 1946–73

The three equations comprising the New Classical model are given by

$$DG_t = \gamma_0 + \gamma_1 DG_{t-1} + \gamma_2 UN_{t-1} + \gamma_3 WAR_t + \epsilon_{1t} \tag{B1}$$

$$DM_t = \beta_0 + \beta_1 DM_{t-1} + \beta_2 DM_{t-2} + \beta_3 UN_{t-1} + \beta_4 E_{t-1}(FEDV_t) + \epsilon_{2t} \tag{B2}$$

where

$$E_{t-1}(FEDV_t) = FEDV_t - 0.8(DG_t - \gamma_0 - \gamma_1 DG_{t-1} - \gamma_2 UN_{t-1} - \gamma_3 WAR_t)$$

and

$$UN_t = \alpha_0 + \alpha_1 MIL_t + \alpha_2 MINW_t + \alpha_3 DMRH_t + \alpha_4 DMRH_{t-1}$$
$$+ \alpha_5 DMRH_{t-2} + \epsilon_{3t} \tag{B3}$$

where

$$DMRH_{t-j} = DM_{t-j} - \beta_0 - \beta_1 DM_{t-j-1} - \beta_2 DM_{t-j-2}$$
$$- \beta_3 UN_{t-j-1} - \beta_4 E_{t-j-1}(FEDV_{t-j}).$$

It is assumed that

$$\begin{bmatrix} \epsilon_{1t} \\ \epsilon_{2t} \\ \epsilon_{3t} \end{bmatrix} \sim NID\,[0,\,V],$$

where

$$V = \begin{bmatrix} \sigma_1^2 & 0 & 0 \\ 0 & \sigma_2^2 & 0 \\ 0 & 0 & \sigma_3^2 \end{bmatrix}.$$

1. Estimation

The most straightforward method of estimating the equations as a system imposing the cross–equation restrictions and the assumption that V is diagonal is to transform the i'th equation by $1/\sigma_i$ and then to stack the equations. Ignoring the cross–equation restrictions for the moment, suppose that equations (B1)–(B3) can be depicted as

$$DG_t = X_{1t}\alpha_1 + \epsilon_{1t}$$

$$DM_t = X_{2t}\alpha_2 + \epsilon_{2t}$$

$$UN_t = X_{3t}\alpha_3 + \epsilon_{3t}.$$

Then the stacked system is given as

$$\begin{bmatrix} DG_t/\sigma_1 \\ DM_t/\sigma_2 \\ UN_t/\sigma_3 \end{bmatrix} = \begin{bmatrix} X_{1t}/\sigma_1 & 0 & 0 \\ 0 & X_{2t}/\sigma_2 & 0 \\ 0 & 0 & X_{3t}/\sigma_3 \end{bmatrix} \begin{bmatrix} \alpha_1 \\ \alpha_2 \\ \alpha_3 \end{bmatrix} + \begin{bmatrix} v_{1t} \\ v_{2t} \\ v_{3t} \end{bmatrix}, \qquad (B4)$$

in which $v_{it} = \epsilon_{it}/\sigma_i$ for i = 1, 2, 3. If there were no cross–equation restrictions, equation (B4) could be estimated by OLS. However, the presence of cross–equation restrictions means we need to use a non–linear estimating procedure for purposes of efficiency. In this paper, the parameters are estimated by the maximum likelihood approach used in the computer package Shazam (see White (1978, 1988) and Byron (1987) for details).

It is necessary to obtain a consistent estimate of σ_i for each equation. It is possible to estimate unrestricted forms of equations (B2) and (B3), but given that there are 28 observations there will be few degrees of freedom (especially in estimating equation (B3)) if this approach is followed. Instead, we have used the three step estimating procedure as in Rush and Waldo (1988) and Pesaran (1988), together with Pagan's (1984) results and those in Theorem 5 of Appendix A.

For σ_1: The estimated error variance from OLS applied to equation (B1) provides a

consistent estimate of σ_1^2. To obtain this estimate, equation (B1) is estimated over the period 1946–73.

For σ_2: The OLS residuals from the equation just estimated, denoted as DGR_t, are used to obtain $E_{t-1}(FEDV_t) = FEDV_t - 0.8DGR_t$. Equation (B2) is then estimated by OLS with this variable over the period 1946–73. The estimated error variance from this OLS regression provides a consistent estimate of σ_2^2 by applying Pagan (1984)'s results.

For σ_3: The OLS residuals from the equation just estimated, denoted as $DMRH_t$, are then used to create $DMRH_{t-1}$ and $DMRH_{t-2}$ (initial observations are set equal to zero). Equation (B3) is then estimated by OLS with these variables over the period 1946–73. By the arguments in Theorem 5 of Appendix A, the estimated error variance from this OLS regression provides a consistent estimate of σ_3^2.

Suppose the above procedure yields estimates $\hat{\sigma}_1$, $\hat{\sigma}_2$ and $\hat{\sigma}_3$. These are used to transform equations (B1), (B2) and (B3), and the resulting system is estimated by maximum likelihood imposing the cross–equation restrictions. The same transformation is used in calculating the RESET, serial correlation and non–nested tests when the New Classical model is the null.

Note: The OLS regressions to obtain $\hat{\sigma}_1$, $\hat{\sigma}_2$ and $\hat{\sigma}_3$ do not correspond to those in Rush and Waldo (1988) and Pesaran (1982, 1988) in either the time periods used or the method of correcting for $E_{t-1}(FEDV_t)$.

2. Testing

2.1 RESET Test

Using the maximum likelihood estimates, let the fitted values from equations (B1) – (B3) be denoted as $\hat{y}_{it}$ (i = 1, 2, 3). The extra regressors involving squared fitted values of each dependent variable, namely

$$\begin{bmatrix} \hat{y}_{1t}^2/\sigma_1 & 0 & 0 \\ 0 & \hat{y}_{2t}^2/\sigma_2 & 0 \\ 0 & 0 & \hat{y}_{3t}^2/\sigma_3 \end{bmatrix}$$

are added to equation (B4) and the system is re-estimated. The appropriate likelihood ratio test is calculated for the system as a $\chi^2(3)$ test statistic (see Ramsey (1969, 1974) for details regarding the single-equation testing procedure). The relevant test statistic for a single equation may be obtained as a $\chi^2(1)$ statistic by adding only one column of the above matrix at a time, such as

$$\begin{bmatrix} \hat{y}_{1t}^2/\sigma_1 \\ 0 \\ 0 \end{bmatrix}$$

in order to test the first equation, assuming that the second and third equations are specified correctly.

2.2 Serial Correlation Test

Using the maximum likelihood estimates, the residuals from each equation, namely $\hat{\epsilon}_{1t}$, $\hat{\epsilon}_{2t}$ and $\hat{\epsilon}_{3t}$ are calculated. The extra regressors involving lagged values of these residuals, namely

$$\begin{bmatrix} \hat{\epsilon}_{1t-1}/\sigma_1 & 0 & 0 \\ 0 & \hat{\epsilon}_{2t-1}/\sigma_2 & 0 \\ 0 & 0 & \hat{\epsilon}_{3t-1}/\sigma_3 \end{bmatrix}$$

are added to equation (B4) and the system is re-estimated (initial values of the lagged residuals are set equal to zero). The appropriate likelihood ratio test is calculated for the system as a $\chi^2(3)$ test statistic (see Breusch and Godfrey (1981) and Godfrey (1978) for

details regarding the single–equation testing procedure). The relevant test statistic for a single equation may be obtained as a $\chi^2(1)$ statistic by adding only one column of the above matrix at a time, such as

$$\begin{bmatrix} \hat{\epsilon}_{1t-1}/\sigma_1 \\ 0 \\ 0 \end{bmatrix}$$

in order to test the first equation, assuming that there is no serial correlation in the second and third equations.

2.3 Heteroskedasticity Test

Using the maximum likelihood estimates, let the fitted values of each equation be denoted as $\hat{y}_{it}$ (i = 1, 2, 3) and the corresponding residuals be $\hat{\epsilon}_{it}$ (i = 1, 2, 3). The Lagrange multiplier (LM) test for heteroskedasticity in equation i is based on TR^2, that is, the sample size times the coefficient of determination from the auxiliary regression of $\hat{\epsilon}_{it}^2$ on an intercept and $\hat{y}_{it}$ for each equation (i = 1, 2, 3). The LM test, which is distributed asymptotically as $\chi^2(1)$ under the null hypothesis, is a test of heteroskedasticity in the particular equation considered, assuming there is no heteroskedasticity in the other two equations (see Pagan and Hall (1983a, b) for further details).

2.4 Normality Test

Based on the maximum likelihood estimates, the residuals for each equation of the New Classical system are obtained. The Lagrange multiplier test of Bera and Jarque (1981), based on the third and fourth moments of the empirical distribution, is a test of normality in the particular equation considered, assuming there is no non–normality in the remaining two equations. The test statistic is asymptotically distributed as $\chi^2(2)$ under the null hypothesis of normality.

2.5 Non—nested Tests: Original New Classical Model as Null

Using the original Keynesian model, namely

$$UN_t = \phi_0 + \phi_1 MIL_t + \phi_2 MINW_t + \phi_3 DM_t + \phi_4 DM_{t-1}$$

$$+ \phi_5 DG_t + \phi_6 t + \phi_7 WAR_t + \epsilon_{4t}, \tag{B5}$$

$$\epsilon_{4t} \sim NID(0, \sigma_4^2) \quad \text{for } t = 1, 2, \cdots, T$$

the non—overlapping variables between equations (B3) and (B5) are DM_t, DM_{t-1}, DG_t, t and WAR_t. Therefore, the variables added to equation (B4) to calculate the asymptotic F test are

$$\begin{bmatrix} 0 & 0 & 0 & 0 & 0 \\ 0 & 0 & 0 & 0 & 0 \\ DM_t/\sigma_3 & DM_{t-1}/\sigma_3 & DG_t/\sigma_3 & t/\sigma_3 & WAR_t/\sigma_3 \end{bmatrix}$$

from which the likelihood ratio test, distributed as $\chi^2(5)$ under the null, may be computed straightforwardly.

Let the OLS fitted value from the original Keynesian model in equation (B5) be denoted by $\tilde{UN}_t$. To obtain the systems version of the J test of Davidson and MacKinnon (1981), add the regressor

$$\begin{bmatrix} 0 \\ 0 \\ \tilde{UN}_t/\sigma_3 \end{bmatrix}$$

to equation (B4) and re—estimate the system by maximum likelihood imposing the cross—equation restrictions. The appropriate likelihood ratio test based on estimating the system is asymptotically distributed as $\chi^2(1)$ when the original New Classical model is the null hypothesis.

Denote the fitted values of UN_t using maximum likelihood estimates of the original

New Classical model by $\hat{UN}_t$. To obtain the systems version of the JA test of Fisher and McAleer (1981), use OLS to estimate the auxiliary regression model given by

$$\hat{UN}_t = X_{4t}\alpha_4 + \text{error}_t$$

namely, the original Keynesian model with UN_t replaced by $\hat{UN}_t$. Denote the fitted values from the auxiliary regression by $X_{4t}\hat{\alpha}_4$, add the regressor

$$\begin{bmatrix} 0 \\ 0 \\ X_{4t}\hat{\alpha}_4/\sigma_3 \end{bmatrix}$$

to equation (B4), and then re—estimate the system by maximum likelihood imposing the cross—equation restrictions. The appropriate likelihood—based JA test statistic is asymptotically distributed as $\chi^2(1)$ when the original New Classical model is the null.

2.6 Non—nested Test: Original Keynesian Model as Null

Treating equations (B1), (B2) and (B5) as a system and assuming that the system has a diagonal covariance matrix and that equation (B5) can be written as

$$UN_t = X_{4t}\alpha_4 + \epsilon_{4t},$$

then the stacked system is

$$\begin{bmatrix} DG_t/\sigma_1 \\ DM_t/\sigma_2 \\ UN_t/\sigma_4 \end{bmatrix} = \begin{bmatrix} X_{1t}/\sigma_1 & 0 & 0 \\ 0 & X_{2t}/\sigma_2 & 0 \\ 0 & 0 & X_{4t}/\sigma_4 \end{bmatrix} \begin{bmatrix} \alpha_1 \\ \alpha_2 \\ \alpha_4 \end{bmatrix} + \begin{bmatrix} v_{1t} \\ v_{2t} \\ v_{4t} \end{bmatrix}, \quad \text{(B6)}$$

in which $v_{it} = \epsilon_{it}/\sigma_i$, $i = 1, 2, 4$. The previously obtained estimates $\hat{\sigma}_1$ and $\hat{\sigma}_2$ are again used. Equation (B5) is estimated by OLS over the period 1946—73 and the estimated error variance is used to obtain a consistent estimate of σ_4^2. For the value of the likelihood function for the restricted model, equation (B6) is estimated by maximum likelihood

imposing the cross—equation restrictions between equations (B1) and (B2). To obtain the likelihood value of the unrestricted model, the variables

$$
\begin{bmatrix}
0 & 0 & 0 \\
\text{DMRH}_t/\sigma_4 & \text{DMRH}_{t-1}/\sigma_4 & \text{DMRH}_{t-2}/\sigma_4
\end{bmatrix}
$$

are added, where these variables are calculated imposing the cross—equation restrictions between them and equations (B1) and (B2). The asymptotic F test is based on the likelihood ratio statistic, which is asymptotically distributed as $\chi^2(3)$ when the original Keynesian model is the null.

The J test may be calculated by adding $\widehat{UN}_t$, the fitted value of UN_t obtained from maximum likelihood estimation of the New Classical model, to equation (B5) and testing the significance of $\widehat{UN}_t$. The t—ratio associated with the OLS estimate of the coefficient of $\widehat{UN}_t$ in this auxiliary regression is asymptotically distributed as $N(0,1)$ under the null Keynesian model.

Denote the OLS fitted value of UN_t from equation (B5) as $\widetilde{UN}_t$, replace the dependent variable in equation (B4) by

$$
\begin{bmatrix}
\text{DG}_t/\sigma_1 \\
\text{DM}_t/\sigma_2 \\
\widetilde{UN}_t/\sigma_3
\end{bmatrix}
$$

and estimate the system by maximum likelihood subject to the cross—equation restrictions. Obtain the fitted values for $\widetilde{UN}_t$ in the system as $X_{3t}\bar{\alpha}_3$ and perform a t—test of the significance of $X_{3t}\bar{\alpha}_3$ when it is included in equation (B5). The t—ratio associated with the OLS estimate of the coefficient of $X_{3t}\bar{\alpha}_3$ in this auxiliary regression is asymptotically distributed as $N(0, 1)$ under the null.

Since there are no cross—equation restrictions to be imposed at the final stage and

the system has a diagonal covariance matrix, there is no gain in efficiency in using maximum likelihood to estimate the auxiliary equations for the J and JA tests as part of a system.

2.7 Testing the Cross–equation Restrictions: Original New Classical Model

The restricted model is given by

$$DG_t = c_0 + c_1 DG_{t-1} + c_2 UN_{t-1} + c_3 WAR_t + \epsilon_{1t}$$

$$DM_t = b_0 + b_1 DM_{t-1} + b_2 DM_{t-2} + b_3 UN_{t-1}$$
$$+ b_4(FEDV_t - 0.8(DG_t - c_0 - c_1 DG_{t-1} - c_2 UN_{t-1} - c_3 WAR_t)) + \epsilon_{2t}$$

$$UN_t = a_0 + a_1 MIL_t + a_2 MINW_t$$
$$+ a_3(DM_t - b_0 - b_1 DM_{t-1} - b_2 DM_{t-2} - b_3 UN_{t-1}$$
$$- b_4(FEDV_t - 0.8(DG_t - c_0 - c_1 DG_{t-1} - c_2 UN_{t-1} - c_3 WAR_t)))$$
$$+ a_4(DM_{t-1} - b_0 - b_1 DM_{t-2} - b_2 DM_{t-3} - b_3 UN_{t-2}$$
$$- b_4(FEDV_{t-1} - 0.8(DG_{t-1} - c_0 - c_1 DG_{t-2} - c_2 UN_{t-2} - c_3 WAR_{t-1})))$$
$$+ a_5(DM_{t-2} - b_0 - b_1 DM_{t-3} - b_2 DM_{t-4} - b_3 UN_{t-3}$$
$$- b_4(FEDV_{t-2} - 0.8(DG_{t-2} - c_0 - c_1 DG_{t-3} - c_2 UN_{t-3} - c_3 WAR_{t-2})))$$
$$+ \epsilon_{3t} \quad (B7)$$

which contains 15 parameters. The unrestricted model is given by

$$DG_t = \gamma_0 + \gamma_1 DG_{t-1} + \gamma_2 UN_{t-1} + \gamma_3 WAR_t + \epsilon_{1t}$$

$$DM_t = \beta_0 + \beta_1 DM_{t-1} + \beta_2 DM_{t-2} + \beta_3 UN_{t-1} + \beta_4 FEDV_t$$
$$+ \beta_5 DG_t + \beta_6 DG_{t-1} + \beta_7 WAR_t + \epsilon_{2t}$$

$$UN_t = \alpha_0 + \alpha_1 MIL_t + \alpha_2 MINW_t + \alpha_3 DM_t + \alpha_4 DM_{t-1} + \alpha_5 DM_{t-2}$$

$$+ \alpha_6 \text{UN}_{t-1} + \alpha_7 \text{FEDV}_t + \alpha_8 \text{DG}_t + \alpha_9 \text{DG}_{t-1} + \alpha_{10} \text{WAR}_t$$

$$+ \alpha_{11} \text{DM}_{t-3} + \alpha_{12} \text{UN}_{t-2} + \alpha_{13} \text{FEDV}_{t-1} + \alpha_{14} \text{DG}_{t-2}$$

$$+ \alpha_{15} \text{WAR}_{t-1} + \alpha_{16} \text{DM}_{t-4} + \alpha_{17} \text{UN}_{t-3} + \alpha_{18} \text{FEDV}_{t-2}$$

$$+ \alpha_{19} \text{DG}_{t-3} + \alpha_{20} \text{WAR}_{t-2} + \epsilon_{3t} \tag{B9}$$

which contains 33 parameters. Therefore, in going from the unrestricted to the restricted model, 18 cross–equation restrictions are being imposed. Given that there are 21 parameters in the unrestricted UN equation and only 28 observations when the equation is estimated over the period 1946–73, the tests of the cross–equation restrictions should be treated with some caution, especially for the shorter sample period.

2.8 Testing the Cross–equation Restrictions: Revised New Classical Model

To equation (B7), add

$$a_6(\text{DG}_{t-1} - c_0 - c_1 \text{DG}_{t-2} - c_2 \text{UN}_{t-2} - c_3 \text{WAR}_{t-1}) + a_7 t$$

to obtain 17 parameters in the restricted model. To equation (B8), add $\alpha_{21} t$ to obtain 34 parameters in the unrestricted model. Therefore, in going from the unrestricted to the restricted model, 17 cross–equation restrictions are being imposed.

2.9 Testing Anticipated Components

For the original New Classical model, test the joint significance of $E_t(\text{DM}_t)$, $E_{t-1}(\text{DM}_{t-1})$ and $E_{t-2}(\text{DM}_{t-2})$ by adding DM_t, DM_{t-1} and DM_{t-2} to the model in equation (B3), in which case the likelihood ratio test is asymptotically distributed as $\chi^2(3)$ under the null hypothesis. In the case of the revised New Classical model, the joint test of the three monetary expectations as well as the fiscal expectation, $E_{t-1}(\text{DG}_{t-1})$, may be performed by adding DM_t, DM_{t-1}, DM_{t-2} and DG_{t-1} to the model and using the likelihood ratio test, which is distributed as $\chi^2(4)$ under the null.

2.10 Testing Unanticipated Components

In contrast to the test of the anticipated components, tests of the unanticipated components examine the joint significance of the monetary shocks, namely $DMRH_t$, $DMRH_{t-1}$ and $DMRH_{t-2}$ for the original New Classical model, which is supplemented by the fiscal shock DGR_{t-1} in the case of the revised New Classical model. The likelihood ratio tests in the two cases are asymptotically distributed as $\chi^2(3)$ and $\chi^2(4)$, respectively, under the appropriate null hypotheses.

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