# BUDGET DEFICITS AND INTEREST RATES: A COINTEGRATION AND CAUSALITY ANALYSIS

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**Abstract –** This paper argues that existing empirical studies of the relationship between budget deficits and interest rates are vitiated by their failure to take into account the non-stationarity, cointegration, and exogeneity features of the data. We investigate the conditions under which the single-equation framework, typically adopted in the literature, can yield valid inference. We then suggest that, in order to deal with the endogeneity problem, one should use the Phillips-Hansen fully modified estimator, which is asymptotically equivalent to full information methods as long as a single cointegrating vector exists. Furthermore, we employ some recently developed testing procedures for both short- and long-run causality in cointegrated systems, and find a positive relationship between budget deficits and interest rates both in the short and in the long run in the post World War II US economy.

**Keywords:** *Budget Deficits, Cointegration, Granger Causality, Long-Run Causality, Phillips-Hansen Fully Modified Estimator, Ricardian Equivalence, Weak Exogeneity*

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## INTRODUCTION

One of the most important issues for the conduct of fiscal policy is the relationship between budget deficits and interest rates. On a theoretical level, it is a moot point whether such a relationship exists at all, and whether it is a short- or long-term one. Two well-known theoretical models imply a positive relationship between budget deficits and interest rates. The first is the static IS-LM model (adopted, e.g., by Cebula and Belton 1992, Evans 1985, 1987b, Viren 1988), according to which an increase in the budget deficit raises disposable income and thus domestic consumption. This leads to an increase in the real demand for monetary assets which, with a fixed real money stock, causes a rise in the short-run nominal interest rate. It should be obvious then that the IS-LM model implies a short-run relationship between budget deficits and interest rates and has no meaningful long-run implications. A long-run link emerges instead from the loanable funds model (used, e.g., by Correia-Nunes and Stemitsiotis 1995, and Hoelscher 1986), which assumes equilibrium of domestic saving and investment flows. This equilibrium condition yields a reduced form expression for the long-term nominal interest rate, in which the budget deficit is one of its exogenous arguments. In this context, an increase in the budget deficit causes a permanent increase in the long-term interest rate, which, in turn, affects the interest-sensitive components of private investment, thus producing a "long-run crowding out" of private investment. Therefore, in the context of the loanable funds model, the long-term interest rate is the channel through which the budget deficit affects the real economy.

An alternative line of thought, associated with the Ricardian equivalence (RE) proposition, leads to the conclusion that there is no relationship between the two variables in question, either in the short or in the long run. If RE holds, households treat future service taxes as an exact offset to government debt (Barro 1974, 1989). Therefore, it makes no difference to an economy whether a given level of government spending is deficit- or tax-financed. In this context, tax smoothing appears to be the optimal policy, provided that "rational Ponzi games" are not feasible (Blanchard and Fisher 1989). In particular, forward-looking agents anticipate the increase in future taxes associated with higher borrowing, and hence respond to an issue of new government debt by increasing their demand for debt by the same amount, thus leaving interest rates unaffected. Consequently, in a regression of short- or long-term, current interest rates on some measure of government debt policy, the coefficient on this variable should be zero if debt neutrality holds.

Numerous empirical studies have tested the RE hypothesis against the alternative of a positive relationship between budget deficits and interest rates. They cover a wide range of alternative data sets and econometric techniques, and provide mixed evidence[4]. As pointed out by Seater (1993), such studies are generally affected by serious drawbacks, as they typically fail to investigate the stationarity, cointegration and exogeneity features of the data. Testing for

[4] Some studies find a significant positive relationship between interest rates and budget deficits (Hoelscher 1986), others a negative one between short-term rates and budget deficits (Zahid 1988), or no significant relationship (Makin 1983, Hoelscher 1986). Finally, Mascaro and Meltzer (1983) report that the stock of government debt is not related to either short- or long-term rates.

cointegration is logically prior to asymptotic inference, since the latter is affected by the presence of unit roots and the cointegration properties of the series involved. For example, standard asymptotic inference in the presence of unit roots and no cointegration, a case usually referred to as "spurious regression," will produce totally misleading inferences. Moreover, the employment of single-equation techniques in a cointegrated framework with non-exogenous regressors imports bias and nuisance parameter dependencies in the limiting distributions of the estimators (Phillips 1991).

The main innovation of this paper is the econometric methodology used, which specifically addresses these issues. We investigate the conditions under which the single-equation framework, almost invariably adopted in the RE literature, can yield valid inference. We then suggest that, in order to deal with the endogeneity problem, one should use the Phillips-Hansen fully modified estimator, which is asymptotically equivalent to full information methods as long as a single cointegrating vector exists (Phillips and Hansen 1990). Finally, we test for both long- and short-run causality links by using some recently developed testing procedures for causality within cointegrated systems (Johansen 1992, Johansen and Juselius 1992, Mosconi and Giannini 1992, and Hall and Wickens 1993). As for the theoretical framework, we are careful to distinguish between short-run interactions, which can be usefully analysed within a simple IS-LM framework, and long-run linkages, which can be more meaningfully discussed in the context of a loanable funds model[5].

The layout of the paper is as follows. Section 2 discusses the econometric issues. Section 3 presents some empirical results for the US, which suggest that there are causality links between budget deficits and interest rates both in the long and in the short run. Section 4 offers some concluding remarks.

## 2. ECONOMETRIC ISSUES

This section clarifies under what conditions single-equation estimation methods are appropriate for the estimation of a cointegrating relationship, and then advocates the use of such a procedure with desirable small sample properties. The concept of *weak exogeneity* (Engle *et al* 1983, Boswjik 1992) is central to determining the relevant conditions. As shown by Johansen (1992) and Johansen and Juselius (1992), the conditions for the legitimacy of estimating the cointegrating vectors b in the context of lower-dimensional system can be expressed in terms of zero restrictions on the rows of the loading matrix c, where $\Pi=cb'$, $\Pi$ being the long-run matrix in the context of a Vector Error Correction Model (VECM – see Johansen 1988, 1991). Such cointegrating restrictions on the rank of $\Pi$, $r(\Pi)=r$, $0<r<n$, where n is the number of variables in the system, impose a constraint on how many of the variables in the VAR can be weakly exogenous. If $r(\Pi)=r$, then at least r rows of $\Pi$ are non-zero. This in turn implies that at least r

[5] Developing a fully-fledged long-run growth model, such as the Solow model, with its implied relationships between the stock of government debt and the stock of reproducible capital, or alternatively between interest rates and measures of debt policy, is beyond the scope of this paper.

rows of c are also non-zero. Consequently, the number of potentially exogenous variables is at most (n-r), i.e., $(n-k) \leq (n-r)$ or $k \geq r$. This result indicates that a necessary (but not sufficient) condition for employing a single-equation dynamic linear regression model, instead of a VAR, in order to estimate b, is $r(\Pi) = 1$. Single equation analysis is then asymptotically equivalent to full system estimation.

A sufficient condition amounts to n-1 variables being weakly exogenous with respect to the unique cointegrating vector. In the absence of such a property, single-equation estimation is asymptotically biased, and the inference is affected by nuisance parameters in the limiting distributions. In fact, "in stationary time series regression, single equation estimation usually leads to a loss of statistical efficiency, as in the seemingly unrelated regression context. But in cointegrated systems the use of single equation techniques imports bias, nuisance parameter dependencies, and loses optimality. As a result, arguments in favor of system methods in cointegrated systems seem more compelling than similar arguments in a classical regression context'' [Phillips (1991, p.302)].

Note that, despite the superconsistency property of the OLS estimators of the cointegrating vector, their statistical significance cannot be decided on the basis of standard asymptotic theory grounds. This is due to the fact that the asymptotic distribution of this estimator depends on a number of factors, such as exogeneity and serial correlation of the regressors. In the absence of weak exogeneity and in the presence of serial correlation, the asymptotic inference procedures are considerably more complicated, since some modifications are required in order to be able to address the problem within the framework of the LAMN family (see Phillips 1991, Phillips and Loretan 1991). As far as the widespread use of conventional single-equation error-correction-models (ECM) is concerned, Phillips (1988) shows that inference based on these models is very close to achieving optimality, but, in general, it falls short of attaining it. Single equation ECM estimation methods are optimal only when the regressors are weakly exogenous with respect to the cointegrating vector.

By contrast, Phillips and Hansen (1990) put forward a "fully modified" (FM) OLS estimator which semi-parametrically corrects for both endogeneity and serial correlation of the regressors, thus yielding mixed normal asymptotics, namely asymptotically standard t-statistics, and fully modified Wald tests for analytic restrictions which follow $\chi^2$ asymptotics. Three alternative kernels can be used in the estimation (see Hansen 1992), where the value of the bandwidth parameter should be chosen as discussed by Andrews (1991). Monte Carlo evidence shows that this estimator outperforms alternative ones in small samples[6]. It is the FM-OLS procedure, therefore, that we use in the empirical analysis (as well as full system methods, in order to check robustness).

In brief, if all the regressors are weakly exogenous and random walks, then the asymptotic distribution of the OLS estimator from a static regression is normal and free of nuisance parameters. If weak exogeneity holds, but the regressors are serially correlated, then the cointegrating vector can be estimated from a

[6] For more details on these econometric issues, see Caporale and Pittis (1999).

conventional ECM formulation. Finally, if weak exogeneity fails then one should either estimate the full model or take the asymptotically equivalent Phillips-Hansen approach. Full system methods are also preferable when the dimension of the cointegration space is greater than one. In fact, when $r(\Pi)=r>1$, it is unclear how to interpret the estimates obtained from a single equation. Inferences, in such a case, should be drawn in the context of a restricted VECM once the cointegration rank has been determined.

## 3. EMPIRICAL RESULTS

As mentioned above, the IS-LM model implies a positive relationship between the nominal short-run interest rate and the budget deficit. Specifically, a ceteris-paribus increase in the budget deficit raises real disposable income and thus real private consumption. This, in turn, increases the real demand for monetary assets. If the real money stock remains constant, an increase in the nominal interest rate is necessary to equilibrate the money market (Evans 1987b). [7]Moreover, the IS-LM model predicts three additional variables as arguments of the reduced form equation for the nominal interest rate. First, real government consumption. A ceteris-paribus increase in the latter raises domestic demand, which then leads to a rise in real money demand, thus bidding up the nominal interest rate. Second, the expected inflation rate. A ceteris-paribus increase of that variable reduces the real interest rate, thus stimulating investment demand. This increases real money demand, thus causing the nominal interest rates to rise. Third, the real money supply. Other things being equal, a higher real money supply will be willingly held if the nominal interest rate declines. These considerations suggest the following reduced form equation for the nominal interest rate:

$$i_t = a_0 + a_1 g_t + a_2 d_t + a_3 m_t + a_4 \pi_t \qquad (1)$$

where $i_t$ is the nominal short-term interest rate, $g_t$ denotes real government spending, $d_t$ is the government deficit to GDP ratio, $m_t$ is the real money stock, and $\pi_t$ stands for the expected inflation rate. All parameters, except $a_3$, are expected to be positive.[8]

Although equation (1) stems from a short-run model, one can also interpret it as a long-run relationship in the context of a loanable funds equilibrium model, according to which the long- rather than the short-term nominal interest rate is determined by the demand for and supply of funds in the economy (Hoelscher 1983, 1986, Cebula and Belton 1992). Such an approach implies a reduced form expression for the long-term interest rate, which is a positive function of the expected inflation rate and the budget deficit. According to this view, the coefficient on government consumption, $a_1$, should be zero. Moreover, the loanable funds approach predicts that two additional variables will enter the reduced form for the long-term nominal interest rate. The first is the short-term real interest rate, which reflects the intentions of investors to switch to the short-run segment of the

[7] See also Evans 1985a,b, 1986, 1987a, 1988, 1989.

[8] Of course, in its original formulation, the IS-LM model includes government spending (as opposed to deficit ratio) as a variable, and it is entirely static. However, it can easily be respecified in terms of the deficit ratio, and also be endowed with dynamics.

loanable funds markets. The second is real GDP growth rate, which may be used as a proxy for the accelerator effects of the business cycle on investment (Correia-Nunes and Stemitsiotis 1995). However, both variables can be excluded from the equation under consideration when the latter describes a cointegration relationship in a uni-dimensional cointegration space. This is because there is convincing evidence that both variables are I(0). Therefore, augmenting (1) by these stationary variables will not affect the cointegration properties of the system under consideration, which implies that these variables can be dropped from (1).

Consequently, the difference between the short- and long-run versions of equation (1) should mainly be reflected in the frequency of the data used, and the maturity of the interest rate series employed. In the short-run version of (1), quarterly data and three-month interest rates should be used since the impact of transitory shocks is of interest. On the other hand, in its long-run version, annual data and yields on long-term bonds are preferable since more emphasis is placed on the fundamental factors. Concerning the statistical techniques employed, the use of a cointegration framework for the long-run version of (1) is theoretically appealing, since cointegration refers to the long-run. In the short-run version of (1), the use of cointegration techniques might seem inappropriate. However, even then the time series properties of the variables of interest may be such that integration/cointegration techniques are to be preferred on the grounds of the appropriate asymptotic theory for hypothesis testing.

Implicit in both, short- and long-run, formulations is the a priori restriction that all the RHS variables are exogenous, which restricts the dimension of the cointegration space to be at most one. The "theoretical" notion of exogeneity implies that all RHS variables are determined outside the system under study, which in turn implies that these variables do not adjust to deviations from the long-run equilibrium. In other words, the interest rate alone should bear the burden of adjustment towards equilibrium. The statistical equivalent of this "exogeneity concept" is the weak exogeneity of the regressors with respect to the long-run parameters, since under weak exogeneity the cointegrating vector does not enter the equations for the regressors. If, on the other hand, the cointegration rank is greater than one, then the interpretation of these long- run relationships is not straightforward. Indeed, in a recent paper Wickens (1996) argues that cointegrating vectors resulting from an unrestricted VAR cannot be given an economic interpretation unless additional structural information is introduced. In the present case, the existence of more than one cointegrating vector, and the implied endogeneity of some regressors, could be an indication of the fact that the long-run structure reflects the behavior of policy-makers over and above the behavior of private agents in the goods and money markets. In other words, the additional long-run relationship could be seen as a "money supply rule" linking money supply and interest rates for a given fiscal path. Alternatively, it could represent a "fiscal policy rule" relating government expenditure to the budget deficit for a given trajectory of money supply and interest rates. Therefore, the exogeneity/endogeneity status of the relevant variables has to be ascertained empirically in order to provide a meaningful interpretation of the results.[9]

[9] Caution is required when interpreting the results, since it can be legitimately argued that weak exogeneity tests are model specific, as they depend on the adopted marginalization (Geweke 1990). In other words, given the

In the remainder of the paper we present some empirical evidence based on US data. After testing for the order of integration of the series of interest (section 3.1), we begin the analysis by imposing the theoretical restrictions stemming from both the short- and the long-run formulations, where all RHS variables are exogenous (section 3.2.) This implies that r(Π) is at most one. It follows then from Granger's Representation Theorem that under cointegration there exists a unique error correction mechanism (ECM) for the interest rate, which is the only variable to "adjust" to deviations from the long-run equilibrium. The statistical significance of the EC term can then be seen as a test for cointegration under exogeneity of all regressors, as there is a one-to-one mapping between the existence of a long-run relationship and of an EC mechanism for the endogenous variable. In section 3.3, we relax the exogeneity assumption and carry out cointegration analysis in a VAR framework. Having determined the cointegration rank, we test for weak and strong exogeneity of the variables involved with respect to the cointegrating vector(s), and analyse short- and long-run causality links. Our analysis, therefore, tests the null hypothesis under consideration, i.e., whether the deficit-to-GDP ratio affects interest rates, both in the short and in the long run.

### 3.1 Unit Root Tests

Our dataset consists of US annual and quarterly series for government consumption in 1980 prices, the ratio of government deficit to GDP, money stock (M1), and the GDP deflator. We also employ series for the three-month Treasury bill rate at quarterly frequency and for yields on long-term government bonds at annual frequency. The data are obtained from the Department of Commerce, *Business Conditions Digest*, and cover the period 1947:1-1994:4. Table 1 contains the results from Augmented Dickey Fuller (ADF) tests, where the lag length was selected so as to ensure that the residuals are white noise processes (column 5). The null hypothesis of a unit root cannot be rejected for any of the series at hand, quarterly or annual. As for the deterministic components of the series, it appears that, under the null hypothesis of a unit root, their autoregressive representation contains both a drift and a deterministic trend in the case of government expenditure and deficits (and possibly the real money supply), but only a drift in the case of interest rates and inflation.

We have also investigated the possibility that the non-rejection of the unit root hypothesis is due to a structural break of the series within the sample, which has caused a shift either of the mean of the series or of the slope of the trend (Perron 1989, 1990; Ouliaris, Park and Phillips 1989; and Rappoport and Reichlin 1989). The alternative hypothesis is, therefore, that the series are stationary around a broken drift (Table 1, columns 6 and 7) or a broken trend (Table 1, columns 8 and 9). The limiting distribution of the ADF statistic for the former and the latter case can be found in Perron (1990) and (1989), respectively, when the break dates are known a priori. For the period under investigation, prime candidates for a

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fact that the true underlying GDP is not known, only the endogeneity (rather than the exogeneity) status of a variable can be unambiguously established. Moreover, considering an "incomplete" system can lead to very misleading inference (see Caporale and Pittis 1997).

structural break are the years 1973 (first oil shock) and 1980 (second oil shock, when the Federal Reserve decided to switch from interest rate targeting to monetary base control). However, the results still support the null hypothesis of a unit root against the alternative of segmented trends.

### 3.2 Cointegration Analysis: Single-Equation Methods

Residual-based cointegration tests, and standard and fully modified OLS estimates of the static equation (1) are reported in Table 2. (Note that the fully modified estimation procedure was implemented under the assumption that some of the regressors contain deterministic trends). Irrespective of the frequency of the data, the null of no cointegration is strongly rejected by both the ADF and Phillips' $Z_t$ (1987) tests. All the estimated coefficients have the sign predicted by the respective theoretical model, and almost all of them appear to be highly significant. This result suggests that most variables enter the cointegrating vector. A possible exception might be the inflation rate, which appears to be insignificant for some values of the bandwidth parameter. On the contrary, the coefficient on deficits is statistically significant whatever value of the bandwidth parameter or type of kernel is chosen, implying that deficits do have a long-run effect on interest rates. However, this inference is conditional on the assumption that all RHS variables are exogenous[10]. This implies that there exists a unique EC equation for the interest rate, which is the issue we investigate next.

Table 3 reports the results from the estimation of EC equations for each of the five variables under examination as well as the t-tests of the null hypothesis that the EC coefficient $\rho$ is equal to zero (see Kremers et al 1992) for the asymptotic distribution of the t-statistic ($t_{ECM}$) under the null hypothesis of no-cointegration). It can be seen that, for both data frequencies, the only significant EC term appears in the interest rate equation. Therefore, single-equation analysis suggests that deficits have long-run effects on the interest rate. It is interesting to note that this evidence is robust to the frequency of the data (annual or quarterly) and to the selection of the nominal interest rate (short- or long-term). In other words, both short- and long-term interest rates seem to adjust to an underlying equilibrium relationship as defined by equation (1). This evidence is easily interpretable in the context of the loanable funds model, which implies a meaningful long-run relationship. It is also consistent with the fiscal policy predictions of the short-run Keynesian analysis.

### 3.3 Cointegration Analysis: System Methods

Next, we estimate a VAR model. Valid statistical inference about the cointegrating rank of a system can only be drawn if the selected model is a reasonable approximation of the underlying data generation process. Hence, the adequacy of the VAR has to be checked by means of information criteria (Hannan 1980, Schwarz 1978) and a broad range of diagnostics, such as tests for residual autocorrelation and conditional normality for the system as a whole (Spanos

[10] Under weak exogeneity of the regressors, the Phillips-Hansen method corrects for serial correlation of the regressors only.

1986)[11]. The information criteria suggest to select a VAR of second and fifth order for annual and quarterly data, respectively; in both cases the misspecification tests corroborate this choice (Table 4). Moreover, given the previous results on the existence of a deterministic trend in the univariate representations of the series, we specify each VAR as containing an unrestricted vector of constants, and a deterministic trend which lies in the cointegration space only.

Johansen's (1992) cointegration tests (Table 5) suggest that there is a single cointegrating vector in both the annual and the quarterly cases. Furthermore, the Maximum Likelihood (ML) estimates are very close to those obtained when using the Phillips-Hansen fully modified OLS method. In addition, LR tests for the hypothesis that the interest rate does not enter the cointegrating vector lead to the rejection of the null at the 0.037 and 0.028 significance levels for annual and quarterly data, respectively. The corresponding p-values for the hypothesis that the deficit does not enter the cointegrating vector are 0.021 and 0.037. Therefore, system analysis appears to confirm the single-equation results.

### 3.4 Exogeneity and Causality Tests

Further evidence on the robustness of the system analysis results, obtained from full information methods, is provided in Table 6. There the weak exogeneity status of each variable is tested by means of LR tests on the rows of the loading matrix c (see Section 2). By adopting the notation $i_t = x_{1t}$, $g_t = x_{2t}$ , $d_t = x_{3t}$, $m_t = x_{4t}$, $\pi_t = x_{5t}$, we can express the hypothesis to be tested as follows:

$$H_0 : c_i = 0 \qquad i=1,2,...,5 \tag{2}$$

i.e., $x_{it}$ is weakly exogenous with respect to the cointegrating vector.

Equivalently, (2) can be thought of as a long-run non-causality test for $x_{it}$ (Mosconi and Giannini 1992, Hall and Wickens 1993). The results clearly indicate that the interest rate is not weakly exogenous, whilst the budget deficit, money supply and inflation rate are. Hence, the single-equation results seem to carry forward to the full information case, with the exception of the exogeneity status of government expenditure, which is rejected in the latter (but not in the former) case.

The evidence presented so far indicates that a rise in the deficit causes an increase in interest rates in the long run, whereas the opposite does not hold. As a final test, we have investigated whether there are any short-run causality effects of the budget deficit on interest rates, and whether the former is *strongly exogenous*. The relevant hypotheses tested are:

$$H_0 : (\gamma_{13})_j = 0 \tag{3a}$$

i.e., $i_t$ is not "short-run'' caused by $d_t$, and

$$H_0: c_3=0, \ (\gamma_{31})_j = (\gamma_{32})_j = (\gamma_{34})_j = (\gamma_{35})_j = 0 \tag{3b}$$

i.e., $d_t$ is strongly exogenous; where j=1,2 for annual and j=1,...,4 for quarterly series.

The results shown in Table 7 indicate the existence of short-run unidirectional causality from budget deficits to interest rates. The absence of causality in the opposite direction points to a strong exogeneity status for the budget deficit.

[11] See also Hall (1991) for a discussion on the two approaches to choosing the VAR specification.

## 4. CONCLUSIONS

This paper has presented some empirical evidence, based on US data, on the relationship between budget deficits and interest rates, which is of central importance to policy makers. Other studies have examined the empirical relevance of the RE hypothesis by testing its implications for interest rates. However, they have not clearly distinguished between short- and long-run considerations, and they have taken a flawed econometric approach by carrying out only single-equation estimation and by not addressing issues such as the stationarity, cointegration, and exogeneity status of the variables.

Although we do not specify an explicit long-run model, we do attempt to rationalize our findings by using an IS-LM and a loanable funds framework to explain short-run dynamics and long-run relationships respectively. The main contribution of this paper, though, is its reliance on recent developments in non-stationary time series, where the interactions between exogeneity and unit roots are taken into account, in order to obtain valid statistical inference. Having analysed the conditions under which the single-equation framework, which is typical of the existing literature, can legitimately be adopted, we have advocated the use, in a unidimensional cointegrating space, of the Phillips-Hansen fully modified OLS procedure. This has the advantage of being valid even in the presence of endogenous and serially correlated regressors, and of producing statistics which follow standard asymptotics. Furthermore, it has highly desirable small-sample properties. We have also employed some recently developed testing procedures for both short- and long-run causality in cointegrated systems.

The empirical analysis indicates the following. Firstly, all variables employed are I(1) and cointegrated, with the dimension of the cointegrating space being equal to one. Secondly, the budget deficit is strongly exogenous with respect to the cointegrating vector, which implies that causality is unidirectional, and it affects interest rates both in the short and in the long run, a result which is clearly inconsistent with debt neutrality.

**TABLE 1**
**Unit Root Tests**

| series | ADF($\ell$) | | ADF($\ell$) Trend | t-value for trend | P-AR | ADF($\ell$) Broken drift | | ADF($\ell$) Broken trend | |
|---|---|---|---|---|---|---|---|---|---|
| | $\ell$ | t-stat | | | | 1973 ($\ell$=.55) | 1980 ($\ell$=.68) | 1973 ($\ell$=.55) | 1980 ($\ell$=.68) |
| | (1) | (2) | (3) | (4) | (5) | (6) | (7) | (8) | (9) |
| | | | | a. Annual series | | | | | |
| $i_t$ | 1 | -2.21 | -2.63 | 1.53 | 0.09 | -2.22 | -2.42 | -2.56 | -2.78 |
| $g_t$ | 1 | -2.03 | -3.66 | 3.18 | 0.20 | -2.02 | -2.06 | -3.74 | -3.56 |
| $d_t$ | 1 | -1.81 | -.3.30 | -2.68 | 0.23 | -1.82 | -1.93 | -3.40 | -3.42 |
| $m_t$ | 2 | -0.58 | -1.59 | -1.47 | 0.15 | -0.52 | -0.81 | -1.38 | -1.85 |
| $\pi_t$ | 4 | -1.07 | -1.06 | 0.36 | 0.10 | -1.12 | -1.28 | -1.13 | -1.30 |
| | | | | b. Quarterly series | | | | | |
| $i_t$ | 7 | -1.84 | -1.73 | 0.73 | 0.25 | -1.85 | -2.13 | -1.71 | -1.90 |
| $g_t$ | 4 | -2.78 | -3.03 | 2.65 | 0.18 | -1.79 | -1.81 | -3.09 | -2.98 |
| $d_t$ | 4 | -2.00 | -3.23 | -2.72 | 0.22 | -2.01 | -1.97 | -3.26 | -3.22 |
| $m_t$ | 7 | -0.46 | -1.88 | -1.84 | 0.51 | -0.39 | -0.52 | -1.66 | -1.93 |
| $\pi_t$ | 8 | -2.01 | -2.20 | 0.90 | -0.12 | -2.04 | -2.23 | -2.17 | -2.36 |

*Notes:* The 5% critical values for annual (A) and quarterly (Q) data are given in the table below. $\ell$ denotes the ratio of pre-break sample size to total sample size.

| | $\ell$ | A | Q | Source |
|---|---|---|---|---|
| ADF | | -2.93 | -2.88 | Fuller (1976) |
| ADF-trend | | -3.50 | -3.43 | Fuller (1976) |
| ADF- broken drift | 0.5 | -3.45 | -3.34 | Perron (1990) |
| | 0.7 | -3.39 | -3.34 | Perron (1990) |
| ADF-broken trend | 0.5 | -3.96 | -3.96 | Perron (1989) |
| | 0.7 | -3.85 | -3.85 | Perron (1989) |

## TABLE 2

**Residual-based cointegration tests and fully-modified estimates of the static regression $i_t=a_0+ a_1 g_t + a_2 d_t + a_3 m_t + a_4 \pi_t + u_t$ where some of the regressions contain deterministic trends**

| Method | M | $a_0$ | $a_1$ | $a_2$ | $a_3$ | $a_4$ | ADF | $z_t$ | $a_2 = 0$ $\chi^1$ (1) |
|---|---|---|---|---|---|---|---|---|---|
| | (1) | (2) | (3) | (4) | (5) | (6) | (7) | (8) | (9) |
| | | | | | a. Annual series | | | | |
| OLS | | 0.160 | 0.010 | 0.00036 | -0.057 | 0.197 | -5.35* | | |
| | | (0.051) | (0.005) | (0.00006) | (0.011) | (0.084) | ($\ell$ =2) | | |
| | 2 | 0.121 | 0.017 | 0.00045 | -0.053 | 0.113 | | -5.13* | 40.96* |
| | | (0.060) | (0.006) | (0.00007) | (0.013) | (0.097) | | | (0.00) |
| PH - Bartlet | 5 | 0.092 | 0.020 | 0.00045 | -0.048 | 0.108 | | -5.21* | 32.69* |
| Kernel | | (0.071) | (0.007) | (0.00008) | (0.015) | (0.110) | | | (0.00) |
| | 10 | 0.081 | 0.017 | 0.00032 | -0.0418 | 0.157 | | -4.98* | 25.10* |
| | | (0.055) | (0.006) | (0.00006) | (0.012) | (0.089) | | | (0.00) |
| | 2 | 0.124 | 0.015 | 0.00043 | -0.054 | 0.160 | | -5.11* | 42.90* |
| | | (0.063) | (0.006) | (0.00006) | (0.013) | (0.091) | | | (0.00) |
| PH - Parzen | 5 | 0.101 | 0.019 | 0.00047 | -0.052 | 0.068 | | -5.25* | 37.68* |
| Kernel | | (0.070) | (0.007) | (0.00007) | (0.015) | (0.107) | | | (0.00) |
| | 10 | 0.079 | 0.021 | 0.00046 | -0.042 | 0.084 | | -4.92* | 27.12* |
| | | (0.075) | (0.075) | (0.00008) | (0.015) | (0.11) | | | (0.00) |
| | | | | | b. Quarterly series | | | | |
| OLS | | -0.055 | 0.030 | 0.00031 | -0.014 | 0.254 | -4.69* | | |
| | | (0.021) | (0.002) | (0.00003) | (0.004) | (0.041) | ($\ell$ =5) | | |
| | 4 | 0.144 | 0.012 | 0.00042 | -0.055 | 0.151 | | -4.65* | 63.57* |
| | | (0.052) | (0.005) | (0.00005) | (0.011) | (0.063) | | | (0.00) |
| PH - Bartlett | 10 | 0.163 | 0.012 | 0.000044 | -0.061 | 0.076 | | -4.59* | 52.09* |
| Kernel | | (0.061) | (0.006) | (0.00006) | (0.013) | (0.077) | | | (0.00) |
| | 15 | 0.198 | 0.011 | 0.00055 | -0.069 | 0.012 | | -4.51* | 53.10* |
| | | (0.064) | (0.006) | (0.00006) | (0.014) | (0.084) | | | (0.00) |
| | 4 | 0.153 | 0.011 | 0.00044 | -0.058 | 0.148 | | -4.60* | 77.67* |
| | | (0.047) | (0.004) | (0.00005) | (0.009) | (0.057) | | | (0.00) |
| PH - Parzen | 10 | 0.133 | 0.014 | 0.00047 | -0.054 | 0.132 | | -4.57* | 48.98* |
| Kernel | | (0.051) | (0.006) | (0.00006) | (0.012) | (0.073) | | | (0.00) |
| | 15 | 9.155 | 0.012 | 0.00049 | -0.061 | 0.079 | | -4.50* | 48.41* |
| | | (0.066) | (0.006) | (0.00006) | (0.013) | (0.080) | | | (0.00) |

*Notes:* 1. PH stands for Phillips-Hansen (1990) semi-parametric procedure.

2. M is the selected truncation lag.

3. $Z_l$ is the Phillips (1987) test for cointegration, which is constructed from the OLS residuals $\hat{u}_t$ and, is based on the "bias corrected" serial correlation coefficient estimate. The bias correction is given by $\hat{\lambda} = \sum_{j=-M}^{M} w(j/M) T^{-1} \sum_{t} \hat{\varepsilon}_{t-f}\, \hat{\varepsilon}_t$ where $\hat{\varepsilon}_t = \hat{u}_t - \hat{\rho}\, \hat{u}_{t-1}$

and the kernel weights w(.) satisfy the same restrictions as in the Phillips-Hansen (1990) procedure.

4. Confidence values: 5% for ADF; and $Z_t$ = -4.49 in the case of four regressors and trending variables. (*Source:* Phillips and Ouliaris1990)).

5. Numbers in parentheses denote standard errors in columns 2-6, and p-values in column 9

**TABLE 3**
**OLS estimation of error correction mechanisms**

$$\Delta y_t = \alpha_0 + \rho \hat{u}_{t-1} + \sum_i \delta_{0i} \Delta y_{t-i} + \sum_i \sum_j \delta_{ij} \Delta X_{it-j} + \nu_t$$

**Diagnostics (p-values)**

| Series | t-test for $H_0$:ρ=0 | AR($\ell$) | Lin | N | H |
|---|---|---|---|---|---|
| | | **a. Annual series** | | | |
| $i_t$ | -3,10* | 0.10 | 0.37 | 0.34 | 0.53 |
| $g_t$ | 1.55 | 0.26 | 0.44 | 0.10 | 0.75 |
| $d_t$ | -0.56 | 0.15 | 0.07 | 0.11 | 0.09 |
| $m_t$ | 0.62 | 0.08 | 0.12 | 0.72 | 0.13 |
| $\pi_t$ | 1.70 | 0.25 | 0.13 | 0.65 | 0.91 |
| | | **b. Quarterly series** | | | |
| $i_t$ | -3,91* | 0.12 | 0.69 | 0.09 | 0.22 |
| $g_t$ | 0.01 | 0.08 | 0.11 | 0.10 | 0.13 |
| $d_t$ | -0.10 | 0.31 | 0.23 | 0.08 | 0.09 |
| $m_t$ | -0.60 | 0.15 | 0.38 | 0.15 | 0.12 |
| $\pi_t$ | 0.20 | 0.17 | 0.83 | 0.07 | 0.08 |

*Notes:* 1. Critical values for $H_o$: ρ= 0 are at 5%; for alternative values of signal-to-noise ratio q: q=0 (-2,02), q=3 (-1,88), q=8(1,80). *Source*: Kremers, Ericsson and Dolado (1992).

2. AR($\ell$) is a Lagrange multiplier test for no residual serial correlation, where $\ell$=1,4 for annual and quarterly series, respectively.

Lin is a reset type test for non linearity.

N is a skewness and kurtosis residual-based test for normality.

H is a reset-type test for heteroskedasticity.

**TABLE 4**
**Order selection for the unrestricted VAR**

| $\ell$ | HQIC | SIC | SAC (P-value) | SN (P-value) |
|---|---|---|---|---|
| | | **a. Annual series** | | |
| 3 | -23.28 | -21.15 | 0.11 | 0.17 |
| 2* | -23.29 | -21.77 | 0.08 | 0.09 |
| 1 | -22.23 | -21.35 | 0.00 | 0.00 |
| | | **b. Quarterly series** | | |
| 7 | -30.53 | -28.62 | 0.19 | 0.22 |
| 6 | -30.58 | -28.96 | 0.13 | 0.18 |
| 5* | -30.61 | -23.19 | 0.09 | 0.11 |
| 4 | -30.28 | -23.15 | 0.02 | 0.05 |

*Notes:* 1. HQIC and SIC stand for Hannan-Quin and Schwartz information criteria, respectively, whereas SAC and SN denote system autocorrelation and system normality, respectively (Spanos 1986).

2. An asterisk (*) indicates the selected lag length.

**TABLE 5**
**a. System cointegration tests**

| | λ- max | | 5% c.v. | Trace | | 5% c.v. |
|---|---|---|---|---|---|---|
| | Annual | Quarterly | | Annual | Quarterly | |
| $r \leq 0$ | 38.85* | 42.82* | 37.5 | 96.15* | 107.4* | 87.3 |
| $r \leq 1$ | 24.86 | 30.44 | 31.5 | 57.31 | 61.57 | 63.0 |
| $r \leq 2$ | 18.9 | 17.81 | 25.5 | 32.44 | 31.13 | 42.4 |
| $r \leq 3$ | 10.02 | 8.07 | 15.0 | 13.54 | 13.33 | 25.3 |
| $r \leq 4$ | 3.51 | 5.25 | 12.3 | 3.51 | 5.25 | 12.3 |

**b. Estimated cointegrating vector**

| | | $i_t$ | $g_t$ | $d_t$ | $m_t$ | $\pi_t$ |
|---|---|---|---|---|---|---|
| a. | Annual | -1.00 | 0.051 | 0.0006 | -0.056 | 0.330 |
| b. | Quarterly | -1.00 | 0.70 | 0.0005 | -0.162 | 0.94 |

*Notes:* 1. Critical values (c.v.) are obtained from Osterwald-Lenum (1992).

2. Each VAR was estimated with an unrestricted vector of constants and deterministic trend, which lies in the cointegration space only.

**TABLE 6**
**LR tests for weak exogeneity of each variable with respect to the cointegrating vector, $H_o$: $c_i = 0$, i = 1, 2,...,5**

| Series | Annual | Quarterly |
|---|---|---|
| $i_t$ | 4.65* | 4.62* |
| | (0.03) | (0.03) |
| $g_t$ | 29.71* | 18.30* |
| | (0.00) | (0.00) |
| $d_t$ | 0.005 | 2.83 |
| | (0.98) | (0.10) |
| $m_t$ | 0.021 | 1.51 |
| | (0.88) | (0.21) |
| $\pi_t$ | 0.68 | 0.66 |
| | (0.40) | (0.41) |

*Note*: Numbers in parentheses are p-values.

**TABLE 7**
**LR tests for the hypothesis that the deficit does not "short-run" cause the interest rate, and/or strong exogeneity of the deficit**

| | $H_o$: $d_t$ does not "short-run" cause $i_t$, $(\gamma_{13})_i = 0$ | $H_o$: $d_t$ is strongly exogenous $c_3 = 0$, $(\gamma_{31})_j = (\gamma_{32})_j = (\gamma_{34})_j = (\gamma_{35})_j = 0$ |
|---|---|---|
| Annual (j=1,2) | 11.21 | 2.98 |
| | (0.00) | (0.56) |
| Quarterly (j= 1,...,4) | 23.76 | 22.28 |
| | (0.00) | (0.17) |

*Note*: Numbers in parentheses are p-values.

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