The Relationship between Budget Deficits and Capital Inflows: Further Econometric Evidence

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This research note examines the impact of federal deficits on U.S. capital inflows. Expanding on the previous work of Bahmani-Oskooee and Payesteh (1994), we employ the relatively new maximum likelihood procedure developed by Johansen (1988) and Johansen and Juselius (1990) to do cointegration tests. The results find a long run relationship between budget deficits and capital inflows. In addition, findings from error-correlation modeling reveal that short-run disequilibria in financial markets are corrected very rapidly, suggesting that these markets are efficient.

I. INTRODUCTION

In a recent issue of this journal Bahmani-Oskooee and Payesteh (1994) have shown that in the United States budget deficits led to capital inflows in the 1973-88 period. Using the Engle-Granger (1987) residual-based two-step cointegration procedure and error-correction model, they found budget deficits and capital inflows to be cointegrated. They attributed the relationship between budget deficits and capital inflows to rising interest rates caused by public borrowing and to improved expectations based on increased economic growth resulting from budget stimulus. The evidence provided by Bahmani-Oskooee and Payesteh (hereafter BP) has made an important contribution to our understanding of the effect of budget deficits on the economy. The purpose of this note is to extend the evidence on the relationship between budget deficits and capital flows by examining the data using the most recent cointegration procedure developed by Johansen (1988) and Johansen and Juselius (1990). Subsequent analysis has suggested that the Engle-Granger cointegration technique suffers from a number of econometric shortcomings. Our results further confirm the existence of a long-run link between budget deficits and capital inflows. Moreover, we find that short-run disequilibria in the relationship of budget deficits

and capital inflows are corrected very rapidly, suggesting that global capital markets are very efficient.

II.   THE METHODOLOGY

The Engle-Granger cointegration procedure (hereafter EG) used by BP suffers from several econometric shortcomings. First, Banerjee, Dolado, Galbraith and Hendry (1993), Davidson and Mackinnon (1993), and Stock (1987) have shown there is considerable small-sample bias in estimates derived from the EG procedure. In addition, Davidson and Mackinnon contend that "a relatively low value of R^2 from the cointegration regression should be taken as a warning that the two-step procedure may not work well" (1993, p. 724). It has been demonstrated by Banerjee, Hendry and Smith (1986) and Banerjee et al. (1993) that the size of the small sample bias is inversely related to the magnitude of R^2 in the EG residual-based cointegrating regression. Since the BP data set consists of only 64 observations, it is vulnerable to this criticism. The cointegration results presented in their Table 2 have low R^2 values (.54 and .32), and BP warn that their results must therefore be treated cautiously (p. 68). The second weakness of the EG procedure is it ignores the possibility of multiple cointegrating relationships. Economic variables can exhibit more than one long-run relationship in a cointegrated equilibrium space. The third weakness of the EG method is that it relies very heavily on a super-convergence result and employs ordinary least-squares estimation (hereafter OLS) to derive the parameter estimates of the long-run or cointegrating equation. However, OLS estimates are extremely sensitive to the arbitrary normalization implicit in the selection of the left-hand side variable of the cointegration regression equation (Enders, 1995; Rao, 1994; Banerjee et al., 1993; Muscatelli and Hurn, 1992]. This suggests that different arbitrary normalizations can yield different empirical outcomes. A fourth difficulty with the EG procedure is it does not incorporate short-run dynamics in the cointegrating regression. Not accommodating short-run dynamics results in increased bias, loss of information and thus reduced efficiency of the parameters of interest in the cointegrating relationships. Finally, and most importantly, the EG procedure does not enable the researcher to test for various restrictions or exclusions on individual elements of the observed cointegrating vectors. In testing of hypotheses related to long-run economic relationships, this drawback of the EG procedure is a serious shortcoming.

The maximum likelihood procedure of Johansen (1988) and Johansen and Juselius (1990)—hereafter J and JJ respectively—is able to overcome the aforesaid shortcomings of the EG procedure. Furthermore, as Gonzalo (1994) has demonstrated in his Monte Carlo study, the Johansen procedure performs better than other estimators of long-run parameters even in the presence of non-normal errors and unknown dynamics. The system-based procedure of J and JJ

provides a natural econometric framework for a combined analysis of the long and short-run behavior of variables of interest.

In the JJ cointegration procedure, two tests, the maximal eigenvalue and trace tests, are used to determine the number of cointegrating vectors (see Enders, 1995; Serletis, 1993; Juselius and Hargreaves, 1992). In the maximal eigenvalue test we test the null hypothesis of r cointegrating vectors against the alternative $r + 1$ cointegrating vectors. In the trace test the null hypothesis is that there are at most r cointegrating vectors against a general alternative (see J, 1988; JJ, 1990). A detailed description and mathematical exposition of the Johansen maximum likelihood procedure can be found in Dickey, Jensen and Thornton (1994), Muscatelli and Hurn (1992), J (1988) and JJ (1990).

III.   THE ESTIMATION AND RESULTS

In exploring the relationship between budget deficits and capital flows, BP measure capital inflow (CAI) as seasonally adjusted net capital flows and budget deficits (BUS) as the seasonally and cyclically adjusted federal budget deficit.[1] They analyze these nominal variables as well as the deflated variables CAIY and BUSY, where CAI and BUS are divided by gross national product.

In this study we report the results for CAIY and BUSY. A number of researchers have suggested that the impact of the budget deficit should be judged relative to the size of the economy (Cebula and Belton, 1993; Cebula and Koch, 1989,1994; Cebula, 1990; Darrat, 1988; Evans, 1985). For this reason, and to facilitate comparison with those studies, we report the results from the ratio specification.[2]

Cointegration procedure requires the pretesting of all the variable-series included in the system for the order of integration. For cointegration, all variables should be integrated of identical order. A variable is said to be integrated of order one, I(1), if its series must be differenced once to become stationary, I(0). In Table 1 we present a battery of unit root tests for stationarity. According to the augmented Dickey-Fuller test (ADF), the null hypothesis of the presence of a unit root cannot be rejected for CAIY and BUSY at the 5 percent significance level. For the first differenced series of these variables, ΔCAIY and ΔBUSY, the unit root null hypothesis is rejected at the 5 percent level of significance [see Dickey and Fuller (1979; 1981)]. The lag lengths in the reported ADF tests are chosen on the basis of residual diagnostics such that the lags are large enough to ensure the residuals in the ADF regressions are white noise.

In Table 1 we also present the results of the Phillips-Perron (1988) test, which allows for a general form of serial dependence and conditional heteroscedasticity and involves essentially a non-parametric adjustment to the Dickey-Fuller statistics. As revealed in Table 1, the Phillips-Perron test results echo the conclusions of the ADF tests with the exception of the variable CAIY in its level. Since the observed diagnostic statistics from the ADF regression for this variable

Table 1.   **Unit Root Tests**

Variables	ADF Test	Phillips-Perron Test	Johansen Test
CAIY	−2.0996(6)	−6.9507(5)[*]	0.7389(6)
ΔCAIY	−4.3009(6)[*]	−22.380(5)[*]	17.6987(6)[*]
BUSY	−2.0830(4)	−3.2600(5)	2.0736(6)
ΔBUSY	−3.9700(4)[*]	−10.273(3)[*]	14.4862(6)[*]

Notes: ADF regressions contain a constant and a time trend. Lags in parentheses. For ADF and Phillips-Perron tests reported here, the 5 percent and 10 per cent significance levels are −3.50 and −3.18, respectively (see Fuller, 1976). The 5 percent rejection region for the Johansen statistic, J, is $\{J \in R \mid > 9.094\}$ (Johansen and Juselius, 1990).
[*]The null hypothesis of the presence of a unit root is rejected at the 5 percent level.

Table 2.   **Dickey-Pantula Sequential Tests**

$$\Delta^3 X_t = \alpha_0 + \alpha_1 \Delta^2 X_{t-1} + \alpha_2 \Delta X_{t-1} + \alpha_3 X_{t-1}$$

Variable	α_1	α_2	α_3
CAIY	−4.5679[*]	−9.0507[*]	−1.2435
BUSY	−7.1064[*]	−5.8522[*]	−1.5439

Note: [*]The null hypothesis $\alpha_1 = 0$ is rejected at the 95% level.
Source: See Dickey and Pantula (1987) and Fuller (1976).

did not reveal non-normality, autocorrelation or heteroscedasticity, we give more credence to the ADF test result for CAIY (Holden and Perman, 1994). Table 1 also presents the results of the J (1988) univariate test for stationarity (Taylor, 1993). The Johansen J-statistics indicate the levels of CAIY and BUSY are nonstationary, but their first differences are stationary.

Dickey and Pantula (1987) demonstrate that the ADF (1981) test can yield incorrect conclusions if more than a single unit root actually exists. To test for the presence of multiple unit roots the Dickey-Pantula test (1987) can be performed. The results are presented in Table 2. The evidence clearly establishes that both the CAIY and BUSY series have a single unit root giving us some comfort in the validity of the ADF test. In sum, the empirical evidence reported in Tables 1 and 2 shows that while CAIY and BUSY are I(1), ΔCAIY and ΔBUSY are I(0), and thus the variables are stationary in first differences.

Table 3 presents cointegration tests based on a vector autoregression (VAR) of the observed variables $x_t' = (CAIY_t, BUSY_t)$. To facilitate comparison with BP (1994), the sample period is 1973I − 1988IV. Details on the data used in this research note can be found in BP (1994). An optimal lag length of one is chosen on the basis of the Schwartz Bayesian Criterion (SBC). The analysis indicated the absence of serial correlation in the residuals of the equation with the chosen lag. The cointegration results are basically robust to increasing lag length of the VAR. The estimates presented in Table 3 show that there is one cointegrating vector between the variables CAIY and BUSY. The actual trace and maximal

Table 3.   **Johansen Cointegration Tests**

	Null Hypothesis	Alternative Hypothesis	Actual Statistic	Critical Values (95%)
I. Maximal Eigenvalue Tests	$r = 0$	$r = 1$	46.4093^*	15.752
	$r \leq 1$	$r = 2$	3.3708	9.094
II. Trace Tests	$r = 0$	$r \geq 1$	49.7801^*	20.168
	$r \leq 1$	$r = 2$	3.3708	9.094

Notes: r is the number of cointegrating vectors. VAR lag length = 1. The critical values are from Johansen and Juselius (1990).
*Denotes a rejection of the null hypothesis of no cointegration at the 5 percent significance level.

Table 4.   **Inferences on the Long-run Relationships**

Test	Test Statistic	Critical Value
Linear Trend [$\chi^2 = 1$]	0.2605	3.8415
β_0^* [$\chi^2 = 1$]	19.5442^{**}	6.6349
β_1^* [$\chi^2 = 1$]	26.3888^{**}	6.6349

Notes: β_0^* and β_1^* are the parameters of the cointegrating equation.
**indicates significant at the 1 percent level.

eigenvalue statistics exceed the critical values, rejecting the null hypothesis of no cointegrating vector at the 95% confidence level. However, both trace and maximal eigenvalue statistics fail to reject the null hypothesis that the number of cointegrating vectors is less than or equal to one at the 95% confidence level. Thus, the results confirm the existence of a unique long-run economic relationship between CAIY and BUSY.

Since a clear upward trend in the data could not be found, following JJ (1990) we performed the likelihood ratio (LR) test to determine if a linear trend exists in the data. As shown in Table 4, the *LR* statistic rejects the hypothesis that there is a linear trend and we conclude that it is appropriate to incorporate a constant in the cointegrating vector.

To impart economic meaning to the cointegrating vector, we normalize the vector by the negative value of the reported CAIY coefficient. Thus, the normalized long-run linear relationship is:

$$CAIY = -0.0209 - 0.9943BUSY \tag{1}$$

Since the budget deficit is defined as government revenue minus government spending, if deficits lead to capital inflow we expect the coefficient of the variable BUSY to be negative. Thus, the sign of the coefficient of BUSY is consistent with the underlying economic hypothesis and its magnitude is slightly larger than −0.77 reported by BP. As shown in Table 4, the *LR* test results reveal that for the constant and the coefficient of BUSY the null hypothesis of no significant

Table 5.   **An Error-Correction Model**

Regressor	Dependent Variable
	ΔCAIY
CONSTANT	0.0002
	(0.1447)
ΔBUSY$_{t-1}$	0.5142
	(2.0066)*
RESID$_{t-1}$	-0.9864
	(-7.8298)**

Other Statistics	
$\overline{R}^2$	0.51
DW	2.16
F	32.46
LM***	11.17
RESET***	2.29
NORM***	2.08
HET***	0.51

Notes: t-values in parentheses. *significant at the 5 percent level. **significant at the 1 percent level. ***observed chi-square values.

difference between the restricted and unrestricted models is rejected at the 99% confidence level. Therefore, the estimated results support the view that in the United States, during the sample period, budget deficits and capital inflow formed a long-run link and budget deficits attracted foreign capital.

It has been shown by EG (1987) that cointegrated series have an error-correction (EC) representation and the error-correction mechanism implies that the variables are cointegrated (Boswijk and Franses, 1992; Arize, 1994). Thus, error-correction modeling provides an alternative test for discerning the long-run equilibrium economic relationship between variables. The vector of lagged residuals, RESIDS$_{t-1}$, from the cointegrating regression is used as the error correction term to explain the short-run dynamics of the hypothesis. A parsimonious dynamic error-correction model is estimated and the results are presented in Table 5. The coefficient of the error-correction term, RESIDS$_{t-1}$, is negative and significantly different from zero at the 1% level. This result further supports the existence of cointegration between CAIY and BUSY. In addition, the diagnostic tests for the presence of non-normality, specification error and heteroscedasticity do not reveal any of these conditions, although the LM test for the presence of serial correlation suggests some possibility of serial correlation. Thus, we conclude that the results are sufficiently robust.

On average, the magnitude of the coefficient of the error-correction term, also called the coefficient of speed of adjustment, indicates that nearly 99% of the change in CAIY can be attributed to the disequilibrium between the actual capital inflow and the long-run steady state capital inflow. Since the magnitude

of the speed of adjustment coefficient is not zero, it implies that Granger causality for cointegrated variables is valid (Enders, 1995, p. 367). The large magnitude of the observed speed of adjustment coefficient also implies that capital markets operate efficiently, aided by the presence of low transaction costs, limited regulation, efficient transmission and communication of information, and the absence of capital controls in the United States financial market. This efficiency undoubtedly grew over the time period studied and has expanded in subsequent years. These results also suggest that capital flows act as a force to mitigate the problem of crowding out that might otherwise result from budget deficits.

The results here can be compared with other studies investigating the impact of budget deficits on key economic variables. Using alternative methodologies these studies have found that budget deficits raise long term interest rates (Cebula, 1990; Hoelscher, 1986; Cebula and Belton, 1993) but not short term rates (Hoelscher, 1983; Ostrosky, 1990; Cebula and Belton, 1993). Moreover, Abell (1990) has found that long term interest rates affect capital inflows, while Cebula and Koch (1989, 1994) and Cebula and Belton (1993) have found that capital inflows dampen interest rate increases.[3] While Evans (1985) and Darrat (1990) found that deficits do not affect long term interest rates, they did not control for capital flows.

In summary, the results presented here are consistent with studies in the literature on budget deficits finding that deficits increase long-term interest rates. Long term interest rate movements, in turn, can attract capital inflows.

IV. SUMMARY AND CONCLUSION

In this empirical research note we extend the previous study of BP who investigated whether a long-run link existed between budget deficits and capital inflows in the United States during the 1973-88 period. In employing econometric techniques developed by J and JJ and error-correction modeling, we find that during the sample period budget deficits and capital inflow were directly related and these variables did not wander, in the long-run, arbitrarily from each other. Furthermore, short-run disequilibria in financial markets that attract foreign capital flows are corrected very rapidly, suggesting that these markets are efficient. These results are consistent with the argument that capital inflows act as a force to mitigate the problem of crowding out. The econometric procedures used in the present study overcome the shortcomings of the procedure BP used in their study and clarify their earlier conclusion that budget deficits generate capital inflows.

NOTES

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1. In this study we ignore the long-running argument regarding whether the actual deficit or the cyclically adjusted deficit is the appropriate measure of fiscal policy. As in BP (1994) we use the latter instead of the former. We note, however, that there is some disagreement in the literature on this point (see Barth et al., 1985; Hoelscher, 1985).

2. The results of the level specification, CAI = F(BUS), did not differ markedly from those of the ratio specification at lags varying from one through five in length.

3. Brazelton (1994) has argued that Cebula and Koch's results are due to market forces as well as tight monetary policy. The findings in this study are based on a longer time period (1973-88) that is not dominated by the contractionary monetary policy of the Volcker period (1979-82), thus placing more weight behind the market forces explanation as opposed to the monetary policy explanation.

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