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ABSTRACT

Smaller Pacific Island countries (PICs), namely Fiji, Samoa, Tonga and Vanuatu, which have independent currencies and central banks, have been experiencing budget deficits for more than two decades. This paper investigates the relationship between budget deficits and inflation in the four PICs by undertaking an empirical study of relationship between budget deficits in the four PICs through a panel econometric analysis. A multivariate framework is adopted with a view to avoiding bias arising out of omission of relevant variables and the methodology employed for estimating a long-run relationship between budget deficits and inflation is the Westerlund error-correction-based panel cointegration test procedure. The study findings confirm the existence of a strong, direct relationship between budget deficits and inflation in all four PICs.

I. Introduction

Beginning from the early 1990s, donors discontinued their aid programmes for supporting recurrent budgets of Pacific island countries (PICs). Donors are now directing their assistance towards assisting specific programs such as education and health, and capital projects such as roads, bridges and ports construction. Governments in PICs, being the largest employers, however are not able to contain their operating budgets. Bonuses and pay revisions have been frequent occurrences, whenever revenues from export taxes and levies are high, especially in boom years.

Because of their failure to build up fiscal space during good years, governments find it to difficult finance budget deficits during lean years when revenues are insufficient to cover the expenditures (Jayaraman 2011). Due to slack tax administration and exercise of discretionary powers of tax exemption by ministers, the gap between revenues and rising expenditures has been widening. Based on experiences in the past during the 1990s, international funding institutions have been unanimous in their criticism of fiscal policies of PICs (ADB 2012, UN ESCAP 2012, IMF 2012a, 2102b, 2012c, 2012d, 2012e) and attribute inflation to annual budget deficits to a larger extent.

This paper seeks to test the hypothesis that budget deficits cause inflation in PICs by undertaking a panel cointegration study covering a thirty one-year period: 1981-2011.

Specifically, the paper examines the long-run relationship by resorting to Westerlund error-correction-based panel cointegration test procedures (Westerlund 2007) and then proceeds to test short-term relationships within an error-correction model.

The paper is organized along the following lines: the second section is a brief review of empirical studies; the third section outlines the modeling procedure adopted in the study; and the fourth section presents the results of the empirical investigation; and the last section presents conclusions with policy implications.

II. A Review of Empirical Studies

Monetary economists are categorical in their criticism that annual budget deficits, if incurred year after year, tend to be funded largely by money creation; and addition to money supply creates excess demand, which leads to rise in price level. Their argument is based on the well-known classical quantity theory of money. However, the empirical evidence on the relationship between budget deficits and inflation is not consistent. In a succinct summary, Habibullah, Cheah and Hamid (2011) note that the results of the studies undertaken during last three decades are mixed. The highlights of their summary are given here.

The earliest studies on the subject were on the United States. Niskanen (1978), Hamburger and Zwick (1981) and Dhakal et al. (1994) came to the conclusion that budget deficits caused inflation. However, studies by Dwyer (1982), Karras (1994), Abizadeh and Yousefi (1998) concluded that there was no connection between budget deficits and inflation. Studies by Aghevli and Khan (1978) on selected developing countries, Chang (1994) on Taiwan and Metin (1998) on Turkey showed that budget deficits led to inflation. While Hondroyiannis and Papapetrou (1997) did not find any direct impact of the budget deficit on inflation in Greece, Darrat (2000) concluded in his study that higher budget deficits had a significant hand in the Greece inflationary process.

Studies led by Cukierman et al. (1992) stressed the importance of central bank autonomy (CBA) in inflation control. Their studies lend support to the view that a high degree of CBA helps mitigate the inflationary bias of policy. Since central banks are not autonomous in most of the developing countries, Brown and Yousefi (1996) came to the conclusion that in the absence of political independence of central banks, monetary policy and price stability were undermined in these countries. Autonomy would then mean that a central bank can refuse to finance government deficits and thus, provide more financial stability than would otherwise possible.

In their study on thirteen Asian developing economies, namely; Bangladesh, India, Indonesia, Malaysia, Myanmar, Nepal, Pakistan, the Philippines, Singapore, South Korea, Sri Lanka, Taiwan and Thailand, Habibullah, Cheah and Hamid (2011) analyzed the relationship between budget deficit and inflation. These authors took into consideration the role of money supply by examining its impact on inflation. By identifying the causal direction among the

three variables, the study provided an additional piece of evidence on the growing body of literature on the budget deficits-money- inflation nexus. In the next section, we proceed to examine the long term relationship between budget deficits and inflation in Fiji.

III. Data, Modeling and Methodology

3. 1 Data

The panel study, which excludes Papua New Guinea, which is an outlier in terms of its large land area and natural resources and diversified export basis, is therefore confined to smaller PICs. Among the smaller PICs, we are constrained to omit Solomon Islands as its data series on government finances are not complete. The paper therefore focuses on four PICs: Fiji, Samoa, Tonga and Vanuatu. The panel empirical investigation employs annual data series on consumer price index and budget deficit as percent of GDP for thirty one-year period: 1981-2011.

As budget deficits tend to result in rise in money supply and consequently lead to depreciate the nominal exchange rate, which further contribute to inflation. Further, with a view to minimizing any bias due to omission of relevant variables, we strengthen the analysis by adding two variables: namely broad money (M2) in nominal prices and exchange rate in nominal terms.

Time series data on consumer price index, broad money and nominal exchange rate were obtained from *International Financial Statistics CD Rom* (2012). Data on budget deficit relative to gross domestic product (GDP) in nominal prices were obtained from *Asia Pacific Developing Countries Key Indicators* (Asian Development 2012). The variables, apart from budget deficit-to-GDP, are transformed into natural logarithm for the analysis throughout the study. Table 1 presents summary statistics of data used in the study.

Table 1. Panel data for study

	Consumer price index (2005=100)				Budget deficit-to-GDP (%)				Broad money M2 (Domestic currency, million)				Exchange Rate (Domestic currency/US$)			
Country	Fiji	Samoa	Tonga	Vanuatu	Fiji	Samoa	Tonga	Vanuatu	Fiji	Samoa	Tonga	Vanuatu	Fiji	Samoa	Tonga	Vanuatu
1981-1985 (average)	40.40	29.84	21.95	42.06	6.09	22.65	-1.00	27.66	439.00	38.12	17.70	7777.40	1.02	1.58	1.11	97.73
1986-1990(average)	53.26	45.20	36.36	58.02	5.40	-2.63	0.89	18.57	748.80	90.68	35.52	14673.60	1.35	2.20	1.35	110.69
1991-1995(average)	71.07	60.37	49.85	75.82	3.12	16.59	0.27	4.67	1377.18	133.80	53.86	24611.00	1.48	2.49	1.32	115.03
1996-2000(average)	82.71	72.88	56.94	85.20	3.80	-1.00	1.10	3.14	1452.14	226.32	81.58	32706.20	1.79	2.85	1.47	124.37
2001-2005(average)	94.88	88.99	83.51	96.72	5.08	1.08	-2.30	0.99	2138.22	389.95	156.96	37211.28	1.96	3.06	2.08	125.55
2006-2011(average)	121.43	126.01	126.61	114.52	2.25	3.74	-2.06	0.29	3745.41	688.46	282.36	57353.16	1.77	2.56	1.92	99.38
Over the whole period (1981-2011)																
Mean	78.11	71.62	63.95	79.48	4.25	6.54	-0.48	8.89	1702.56	270.56	108.67	29587.68	1.57	2.47	1.56	112.08
Standard deviation	27.61	32.38	35.79	24.64	2.21	10.89	2.74	11.33	1143.80	231.36	94.82	16578.91	0.37	0.55	0.38	13.64
Maximum	137.42	137.68	138.91	120.21	8.67	37.13	7.97	38.82	4347.24	754.99	299.20	60145.70	2.28	3.48	2.20	145.31
Minimum	35.62	21.99	18.39	38.58	-0.49	-8.84	-6.02	-2.91	364.00	23.10	13.40	3641.00	0.88	1.03	0.87	87.83

Source: Authors' calculations

3.2 Model

The model employed for the study is written as follows:

(1) $\ln P_{\mathrm{it}} = \beta_{i0} + \beta_1 bdr_{\mathrm{it}} + \beta_2 \ln M2_{\mathrm{it}} + \beta_3 \ln er_{\mathrm{it}} + \beta_4 StrBr_{it} + \varepsilon_{it}$

where,
P is consumer price index (2005 = 100);
*M*2 is broad money (in local currencies) in index numbers;
bdr is ratio of budget deficit-to-GDP (percent);
er is exchange rate (domestic currency per US dollar) in index numbers;
StrBr is an artificial variable to capture the effects of structural breaks, which are observed from estimated errors and based on Wald tests on parameter restrictions.
StrBr = 1 for years when structural breaks are observed, and *StrBr* = 0 for the other years.

The hypotheses to be tested are: (i) *bdr* and ln*P* are directly associated; hence sign of the estimated coefficient *bdr* should be positive; (ii) ln*M*2 and ln*P* are directly associated; hence sign of *M*2 should be positive; and (iii) ln*er* and ln*P* are directly associated; hence sign of *er* should be positive.

3.3 Panel Unit Root Test

Before undertaking the econometric analysis, the first critical step is to verify the order of integration of each of the time series of the variables concerned. Accordingly we resort to test the null hypothesis of panel series being non-stationary by employing Breitung test, which assumes the data are generated by an AR(1) process. For higher-order processes, the first-differenced and lagged-level data are replaced by the residuals from regressions of those two series on the first lags of the first-differenced data. A number of lags are therefore be used to remove higher-order autoregressive components of the series.
The Breitung testing procedure follows four steps:

1) We run $\Delta V_{it} = \sum_{p=1}^{P_i} \vartheta_{ip} \Delta V_{it-p} + e_{it}$ and obtain the residuals $\hat{e}_{it}$.
2) Run $V_{it-1} = \sum_{l=1}^{L_i} \varphi_{ip} \Delta V_{it-l} + v_{it-1}$ and obtain the residuals $\hat{v}_{it-1}$.
3) Apply forward orthogonalization transformation to $\hat{e}_{it}$ and $\hat{v}_{it-1}$ to obtain e^*_{it} and v^*_{it-1}.
4) Run $e^*_{it} = \rho v^*_{it-1} + \varepsilon^*_{it}$, where the error term is asymptotically distributed.

In the above procedure, Δ is the first difference operator and V_i is one panel series. The lag order *P* is allowed to vary across cross-sections. The null hypothesis $\rho = 1$ suggests that the panel series contains a unit root. The alternative hypothesis $\rho < 1$ suggests that the panel series is stationary. If V_i is non-stationary, we test for unit root of first difference of V_i, and V_i is said to be integrated of order one, i.e. I(1) if ΔV_i becomes stationary.

3.4 Westerlund Error-Correction-Based Panel Cointegration Test

If all variables are found integrated of order one, the next step is to investigate whether the

variables share a cointegrating relationship. We apply error-correction based Westerlund panel cointegration tests (2007) to serve the above purpose. Westerlund (2007) developed cointegration tests which have good small-sample properties with small size distortions and high power relative to other popular residual-based panel cointegration tests. The underlying idea is to test for the absence of cointegration by determining whether the individual panel members are error correcting. The error-correct model (ECM) for the current analysis assumes the following form:

$$(3)\ \Delta \ln(P)_{it} = \phi_i + \sum_{p=1}^{P_i} \varsigma_{ip} \Delta \ln(P_{it-p}) + \sum_{k=1}^{K} \sum_{p=0}^{P_i} \xi_{k,ip} \Delta X_{k,it-p} + \alpha_i \hat{\varepsilon}_{it-1} + e_{it}$$

where α_i provides an estimate of the speed of error-correction towards the long-run equilibrium. The maximum number of lags *p* can be determined by using Akaike information criterion, Schwarz critierion or Hannan-Quinn criterion. Long-run equilibrium between ln(*P*) and *X*s will be evidenced by a negative coefficient of the error correction term $\hat{\varepsilon}_{i,t-1}$.

Four test statistics are developed to provide critical values for deciding the existence of panel cointegrating relationship. The group-mean statistics Gα and Gτ test statistics test H_0: $\eta_i = 0$ for all *i* versus H_1: $\eta_i < 0$ for at least one *i*. These statistics start from a weighted average of the individually estimated η_i's and their t-ratios, respectively. Therefore rejection of H_0 is taken as evidence of cointegration for at least one of the cross-sectional units. The panel statistics Pα and Pτ are developed to allow both the parameters and dimension of Equation (3) to differ among *i*. The Pα and Pτ statistics test H_0: $\eta_i = 0$ for all *i* versus H_1: $\eta_i < 0$ for all *i*. Rejection of H_0 is therefore taken as evidence of cointegration for the panel as a whole. The above tests allow for an almost completely heterogeneous specification of both the long- and short-run parts of the error-correction model. Furthermore, if the cross-sectional units are suspected of being correlated, robust critical values can be obtained through bootstrapping.

IV. Empirical Results

4.1 Breitung Panel Unit Root Test

For ascertaining the order of integration of the variables in our model, we applied the Breitung panel unit root test, testing the null hypothesis of non-stationarity. The test statistics are summarized in Table 2. The Breitung test statistics for levels of series under consideration are found smaller than the 5 per cent critical value. However, when we subject the first difference of these variables to the Breitung test, we find the test statistics exceed the 5 per cent level critical value, leading us to conclude that all panel series described in the above are each integrated to order one, i.e. $I(1)$.

Table 2 Results of Breitung Panel Unit Root Tests

	Level					First difference				
	Trend	*Panel means*	*# lags*	*λ-stat*	*p*-value	*Trend*	*Panel means*	*# lags*	*λ-stat*	*p*-value
ln*PI*	No	Yes	2	2.2504	0.9878	No	No	2	-2.1468	0.0159
bd/*GDP*	No	Yes	2	-0.9631	0.1678	No	No	2	-6.4823	0.0000

ln*M*2	Yes	Yes	2	0.3729	0.6454	No	No	2	-2.9667	0.0015
ln*er*	No	Yes	2	-0.8633	0.1940	No	No	2	-4.8038	0.0000

4.2 Westerlund ECM Panel Cointegration

Since all panel data series are of I(1), we proceed to testing whether there is any long-run relationship between price index (*P*) and budget deficit (*bdr*). Panel cointegration test results by employing the Westerlund (2007) procedures are summarized in Table 3. The Westerlund error-correction-based panel cointegration tests include the follows settings: deterministic constant; automatic lag selection based on Akaike information criterion (AIC) with average lags of 1, 1 lead in the error correction model based on AIC; and the Bartlett kernel window with bandwidth of 3 in the semi-parametric estimation of long-run variances of α_i in Equation (3).

Table 3: Westerlund ECM Panel Cointegration Tests

Test	Hypotheses	Statistic	Z-value	p-value
Group-mean statistic Gα	H_0: $\eta_i = 0$ for all *i* H_1: $\eta_i < 0$ for at least one *i*	-18.351	-2.099	0.018
Group-mean statistic Gτ	H_0: $\eta_i = 0$ for all *i* H_1: $\eta_i < 0$ for at least one *i*	-3.642	-2.989	0.001
Panel statistic Pα	H_0: $\eta_i = 0$ for all *i* H_1: $\eta_i < 0$ for all *i*	-19.413	-3.646	0.000
Panel statistic Pτ	H_0: $\eta_i = 0$ for all *i* H_1: $\eta_i < 0$ for all *i*	-3.677	0.155	0.562

Westerlund panel cointegration test statistics include group-mean statistic Gα (-18.351), group-mean statistic Gτ (-3.642), panel statistic Pα (-19.413) and panel statistic Pτ (-3.677). The probability values of Gα, Gτ and Pα are less than 5%, providing evidence of a long-run relationship between price and controlling factors, including budget deficit.

4.3 Panel Long-run effects and short-run disequilibrium

Given the cointegrating relationship between price and controlling factors, estimation of Equation (1) yields non-spurious long-run effects of controlling variables on price level in four PICs. Due to the existence of autocorrelation in the panel residuals, Equation (1) is developed into an autoregressive model as follows:

$$(4)\ \ln P_{it} = \beta_{i0} + \beta_1 bdr_{it} + \beta_2 \ln M2_{it} + \beta_3 \ln er_{it} + \beta_4 StrBr_{it} + \beta_5 \ln P_{it-1} + \varepsilon_{it}$$

Equation (4) is estimated by fixed-effect (within) estimator (FE), random-effects generalized least squares estimator (RE), and dynamic panel-data estimator (DPD). RE estimates are found biased since the null hypothesis of no random effects is not rejected by Breusch and Pagan Lagrangian multiplier test with $\bar{\chi}^2 = 0.00$ and p-value = 1.00. FE estimates are found robust, firstly due to the null of no fixed effects is rejected at the 1% significance level with *F-stat* = 20.52 and *p*-value = 0.00, and secondly that FE estimates are highly consistent with DPD

estimates. Furthermore, the Hausman test, which tests the null hypothesis that the RE estimates are the same as the consistent FE estimates, yields Chi-sq statistic = 59.03 and p-value = 0.00. Therefore, relative to the RE estimator, the FE estimator is more appropriate in the current analysis.

The diagnostic tests for FE estimation of Equation (4) are summarized in Table 4. The test results show that the FE estimation is free from autocorrelation, multicollinearity and non-normality problems. Yet, estimation suffers from the heteroskedasticity problem according to the modified Wald test for groupwise heteroskedasticity. Correspondingly, standard errors are adjusted for four clusters in cross-sectional units in order to obtain unbiased standard errors of estimates.

Table 4: Diagnostic Test Results

Test	The Null Hypothesis	Test statistic	*p*-value
Modified Wald test for groupwise heteroskedasticity	Constant variance of error	$chi^2(4) = 26.63$	0.0000
Breusch-Godfrey LM test of independence	No serial correlation in the error	$chi^2(6) = 5.329$	0.5023
Skewness/Kurtosis joint test	Panel residuals are normally distributed	$chi^2(2) = 1.01$	0.6023
Mean Variance Inflation Factor (VIF)	No multicollinearity if VIF is less than 5	Mean VIF = 3.13	

The FE estimates of Equation (4) with robust standard errors are reported as follows:

$$\ln \hat{P}_{it} = 0.310 + 0.002 bdr_{it} + 0.177 \ln M2_{it} + 0.025 \ln er_{it} + 0.066 StrBr_{it} + 0.642 \ln P_{it-1}$$
$$se = (0.039)(0.0004) \quad (0.026) \quad (0.020) \quad (0.010) \quad (0.050)$$
$$t = (8.01) \quad (3.46) \quad (6.80) \quad (1.26) \quad (6.84) \quad (9.80)$$
$$\#\text{years} = 30 \qquad \#\text{countries} = 4$$
$$R^2\text{-within} = 0.9955 \quad R^2\text{-between} = 0.8044 \quad R^2\text{-overall} = 0.6416$$

The above results reveal there is strong evidence that budget deficit has a positive effect on price level in the four PICs. Specifically holding other factors constant, it is seen that a ten percentage points' rise in budget deficit increases price index by 0.02 percent. We also find that broad money has positive effect on price. The coefficient of 0.177 suggests that a ten percent increase in broad money is associated with 1.77 percent increase in price. The nominal exchange rate, which has the expected positive sign indicating a direct relationship with price level, is found statistically not significant. Since keeping the insignificant exchange rate in the regression does not affect the overall performance of the regression, exchange rate remains in the above equation. The artificial variable *StrBr*, representing years when major structural breaks occurred, has a positive effect on price.[1] The coefficient of 0.066 suggests that, keeping other factors unchanged, major structural break on average is associated with 6.6 percent increase in price index. All explanatory variables, except ln*er*, are found to be highly significant at the 1% level.

Short-run disequilibrium is further assessed from the error correction model in Equation (3), and

[1] The dummy variables *StrBr* is generated based on outliers in estimated residuals and Wald parameter tests.

the FE estimation yields results as follows:[2]

$$\Delta \ln \hat{P}_{it} = 0.04 + 0.001\Delta bdr_{it-1} + 0.10\Delta \ln M2_{it} + 0.09\Delta \ln er_{it} + 0.02StrBr_{it} - 0.14\hat{\varepsilon}_{it-1}$$
$$t = (8.29)\ (2.29) \qquad (2.38) \qquad (2.16) \qquad (1.79) \qquad (-1.99)$$
$$R^2 - within = 0.1591 \quad R^2 - between = 0.6926 \quad R^2 - overall = 0.1753$$

The error correction term is significant at the 5% level and has a negative coefficient: – 0.14, suggesting that on an average 14 percent of disequilibrium will be corrected within one year. In the other words, it takes more than 7 years for the disequilibrium to be corrected, which is a slow adjustment.

V. Conclusions and policy recommendations

The paper undertook an empirical study on the impact of budget deficits on inflation in four PICS, namely Fiji, Samoa, Tonga and Vanuatu, during a thirty-one year-period (1981-2011). The study utilized time series of data on price index, budget deficits, money supply, nominal exchange and an artificial variable for capturing the adverse effects of structural breaks on inflation. The study findings confirm that budget deficits were responsible for rise in price level, just as rise in money supply.

The policy recommendations are straightforward. Authorities have to be aware of the grave implications of budget deficits in terms of the inflationary potential of rise in money supply. Aside from cutting non-essential expenditures to trim the budget making process, policy makers in PICs have to pay greater attention to step up revenue collection and discontinuing the needless incentive schemes of tax exemptions and concessions offered by way of subsidies and discretionary measures by ministers. Since the relationship between money supply and budget deficits is clearly established, central banks have to keep a watch on money supply and advise the governments as far as their autonomy under the existing legislation that would permit.

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[2] The ECM regression is free from problems based on diagnostic tests. The FE estimator is found more appropriate than the RE estimator based on the Hausman test. Test statistics are not presented here to converse space, and are available upon request.

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