Fiscal sustainability in a panel of Asian countries

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Panel cointegration techniques indicate that government revenue (REV) and expenditure (EXP) in a panel of five Asian economies for the period 1974 to 2001 were nonstationary and cointegrated series. However, the cointegration coefficient was significantly less than unity, indicating ‘weak’ fiscal sustainability and the likelihood that policy measures would be needed to put the public finances on a more sustainable basis.

I. Introduction

A common technique for evaluating fiscal policy sustainability is to examine past data to establish whether there is cointegration between government revenue (REV) and expenditure (EXP). Most of the related empirical research has been with respect to the experience of developed economies, whereas in practice, fiscal sustainability problems most often have been confronted by low- and middle-income economies.1

In addition, the unit root and cointegration tests generally employed to establish cointegration between revenue and expenditure typically are tests that have low power in small samples and frequently do not reject the null of a unit (Banerjee, 1999) and frequently give conflicting results. In this article, we exploit panel unit root and cointegration techniques to overcome the problems posed by small samples and employ a dynamic ordinary least squares (DOLS) model to estimate the long-run cointegrating relation between government revenue and expenditure to distinguish between ‘strong’ and ‘weak’ sustainability. In addition, we switch the focus of fiscal sustainability from developed economies to the experience of five developing Asian economies for which consistent data are available. Our results indicate that only the weak sustainability condition is satisfied for the panel of Asian countries, which would seem to justify the steps taken in the early part of this decade by many of them to strengthen their fiscal balances.

II. The Data, Methodology and Results

We use annual data for government revenue and expenditure (in relation to gross domestic product (GDP)) from the International Monetary Fund’s Government Finance Statistics database, which remains the best source of internationally comparable fiscal data. This source allows us to construct a balanced panel for five Asian countries – India, Pakistan, the Philippines, Sri Lanka and Thailand – for the period 1974–2001. The series for government, revenue, expenditure and the overall balance are displayed in Fig. 1 (in percentage of GDP) and indicate some diversity in experience, with persistent fiscal deficits in India, the Philippines (except for the 1994–1997) and Sri Lanka; a persistent surplus in Pakistan until the end of the sample period and a somewhat more varied history for Thailand.

The first step is to analyse the time series properties of the data to determine the persistence of the pooled revenue and expenditure series. We use the panel unit root tests proposed by Levin et al. (2002) and Im et al. (2003) to test the stationarity of the government revenue and expenditure, and the multivariate residual
panel techniques pioneered by Pedroni (2004) to test for cointegration between the series. These tests are constructed by averaging individual augmented Dickey–Fuller (ADF) (Dickey and Fuller, 1979) \( t \)-statistics across cross-section units and their most valuable feature is the degree of homogeneity that they allow. The Levin et al. (2002) test allows for heterogeneity of the intercepts across members of the panel and tests of the null hypothesis that each individual time series in the panel is integrated versus the alternative hypothesis that all individual time series are stationary. It is based on the following pooled ADF equation:

\[
\Delta y_{it} = \alpha_{0it} + \alpha_{1it} + \delta y_{it-1} + \sum_{L=1}^{p_i} \beta_i \Delta y_{it-L} + \varepsilon_{it}
\]  

(1)

where a common \( \delta = \rho -1 \) is assumed, \( \alpha_{0i} \) and \( \alpha_{1i} \) represent country-specific fixed effects and deterministic trends, respectively, and \( p_i \) is the required country-specific degree of lag augmentation to make the residuals white noise that is determined by the conventional step-down procedure. The null of \( H_0: \delta = 0 \) under the assumption that \( \delta_i = \delta \) for all \( i \) is tested against the alternative hypothesis, \( H_1: \delta < \delta_i \) for all \( i \). The test is based on a technique that removes autocorrelation as well as deterministic components. The Im et al. (2003) test allows for heterogeneity in intercepts as well as in the slope coefficients and the test takes the form

\[
\Delta y_{it} = \alpha_{0it} + \alpha_{1it} + \theta_i + \delta y_{it-1} + \sum_{j=1}^{p_i} \beta_j \Delta y_{it-j} + \varepsilon_{it}
\]

(2)

where \( p_i, \alpha_{0i}, \) and \( \alpha_{1i} \) are as in Equation 1 and \( \theta_i \) represents the time dummies used to account for cross-sectional correlation that could result from common shocks affecting all countries in the panel in the period. The null of \( H_0: \delta_i = 0 \) for all \( i \) (i.e. all series have a unit root) is tested against the alternative \( H_1: \delta_i < 0 \) for \( i = 1, 2, \ldots, N_1 \) and \( \delta_i = 0 \) for \( i = N_1 + 1, N_1 + 2, \ldots, N \).

On the assumption that the \( N \) cross-section units are independently distributed, the \( t \)-statistic can be computed as an average of the individual ADF \( t \)-statistics such that

\[
\bar{t}_{NT}(p_i) = \left( \frac{\sum_{i=1}^{N} t_{IT_i}(p_i)}{N} \right)
\]  

(3)
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where \( t_{lT}(p,\rho) \) is the \( t \)-statistic for testing \( \delta_l = 0 \) in each individual ADF regression. In a further step, the above \( t \)-bar statistic is standardized so that it converges to a standard normal distribution, as \( N \) increases. A key strength of this test is that \( \delta_l \) is allowed to differ across countries and only a fraction of panel members is required to be stationary under the alternative hypothesis. The results of the two unit tests are reported in Table 1 and indicate that the null of a unit root in the revenue and expenditure series cannot be rejected by either test for the panel of five Asian countries. Therefore, we can implement a test for panel cointegration between government revenue and expenditure.

The next step is to test for cointegration between government revenue and expenditure in the panel. The available techniques for panel cointegration tests are in essence an application of the Engle and Granger (1987) cointegration analysis. For this study, we follow Kao (1999) and Pedroni (2004), who propose a range of statistics that can be used to determine the presence of cointegration in heterogeneous panels. In each case, the \( t \)-statistics are constructed using the residuals from the following hypothesized cointegrating equation:

\[
REV_t = \alpha + \beta EXP_t + e_t
\]

where \( \alpha \) allows the cointegrating regression to include country-specific fixed effects, \( \beta \) is the long-run coefficient and \( e \) is a residual term. Kao (1999) tests the residuals \( \hat{e}_t \) of the ordinary least squares (OLS) panel estimation by applying DF- and ADF-type tests:

\[
\hat{e}_t = \hat{e}_t \rho_{t-1} + \nu_t
\]

and

\[
\hat{e}_t = \hat{e}_t \rho_{t-1} + \sum_{i=1}^{p} \phi_i \Delta \hat{e}_{t-i} + \nu_t
\]

The null hypothesis of no cointegration, \( H_0: \rho = 1 \), is tested against the alternative hypothesis of stationary residuals, \( H_1: \rho < 1 \). Kao presents five DF and ADF types of cointegration tests in the data panel, the asymptotic distributions of which converge to a standard normal distribution \( N(0,1) \). The \( t \)-statistics are \( DF_{\rho0}^*, DF_{\rho1}^* \) and \( ADF \), which are for cointegration with the endogenous regressors and \( DF_{\rho0}^* \) and \( DF \) are based on assuming strict endogeneity of the regressors. Pedroni (1999) suggests a Phillips–Perron-type panel cointegration test, which implies less strict assumptions with respect to the distribution of the error terms than do the DF and ADF tests described above. He provides two \( t \)-statistics, \( PC_1 \) and \( PC_2 \), which converge to a standard normal distribution. First, under the null hypothesis of no cointegration, the panel autoregressive coefficient estimator \( \hat{\rho}_{NT} \) can be constructed as follows:

\[
\hat{\rho}_{NT} = \frac{\sum_{i=1}^{N} \sum_{t=2}^{T} (\hat{e}_{it-1} - \hat{\lambda}_i) (\hat{e}_{it-1})}{\sum_{i=1}^{N} \sum_{t=2}^{T} \hat{e}_{it-1}^2}
\]

(7)

where \( \hat{\lambda}_i \) acts as a scalar equivalent to the correlation matrix, \( \Gamma \), and corrects for any correlation effect. Pedroni provides the limiting distribution of two \( t \)-statistics:

\[
PC_1 = \frac{T \sqrt{N} (\hat{\rho}_{NT} - 1)}{\sqrt{2}} \Rightarrow N(0,1)
\]

(8)

and

\[
PC_2 = \frac{\sqrt{NT(T-1)} (\hat{\rho}_{NT} - 1)}{\sqrt{2}} \Rightarrow N(0,1)
\]

(9)

The panel cointegration results are reported in Table 2. They indicate that the null hypothesis of no cointegration between government revenue and expenditure can be rejected at conventional significance levels in all cases.

Finally, as the null hypothesis of no cointegration between the series is rejected, the coefficient of the long-run relation between government revenue and expenditure can be estimated. Quintos (1995) distinguishes between a strong and a weak definition of fiscal sustainability. Assuming that government revenues and expenditure are both \( l(1) \), strong solvency occurs if there is cointegration and the slope coefficient \( \beta \) is unity. Weak solvency occurs when \( \beta \) is less than unity. In the context of revenue and expenditure ratios to GDP, Hakkio and Rush (1991) and Cipollini (2001) argue that only the strong condition is appropriate to assess fiscal sustainability as the

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**Table 1. Panel unit root tests for government revenue and expenditure, 1974–2001**

<table>
<thead>
<tr>
<th></th>
<th>Revenue</th>
<th>Expenditure</th>
</tr>
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<tbody>
<tr>
<td>Levin et al. (2002)</td>
<td>0.0631 (0.5252)</td>
<td>-1.6232 (0.9477)</td>
</tr>
<tr>
<td>Im et al. (2003)</td>
<td>2.9962 (0.9986)</td>
<td>-3.6344 (0.9999)</td>
</tr>
</tbody>
</table>

**Notes:** Government revenue and expenditure are in percentage of GDP. The null hypothesis in all tests is that all individual series are nonstationary. Estimates include individual effects and linear trends with lags determined according to the Modified Schwarz Information Criterion.
weak condition may be satisfied but the government still could have difficulty financing a fiscal deficit if the expenditure-to-GDP ratio continuously exceeds the revenue-to-GDP ratio. We estimate the long-run coefficient using the DOLS developed by Kao and Chiang (2001), which is designed for nonstationary panels and corrects the standard pooled OLS for serial correlation and endogeneity of regressors that are normally present in long-run economic relationships. The DOLS procedure involves running the following regression:

\[
REV_{it} = \alpha_i + \beta_i EXP_{it} + \sum_{j=q}^{q} c_{ij} \Delta EXP_{it+j} + \varepsilon_{it}
\]

(10)

where \( t = 1, \ldots, T \) and \( i = 1, \ldots, N \). Equation 10 includes the leads and lags of \( \Delta EXP_{jt} \) in the cointegrating regressions to produce asymptotically unbiased estimators. The estimates of the long-run coefficient are reported in Table 3. We also test whether the cointegration coefficient \( \beta \) is significantly different from 0 and insignificantly different from 1. The estimated \( \beta \) is 0.61, which is far from unity, and both the null hypotheses of \( \beta = 0 \) and \( \beta = 1 \) are easily rejected at the conventional significance levels. Thus, while government revenue and expenditure are nonstationary and cointegrated in the panel of five Asian economies, they appear to have been on only a weakly sustainable path during 1974–2001. The conclusion of weak fiscal sustainability is subject to two important caveats. First, because of weaknesses in the coverage of the fiscal data, the analysis is confined to central government fiscal balances, and developments in the rest of the public sector may well have been such as to provide a stronger record of fiscal performance for the public sector as a whole and perhaps one that would be consistent with ‘strong’ fiscal sustainability. On the other hand, the coverage of the fiscal data also excludes contingent fiscal liabilities that, if realized, will put strain on the public finances.

\[ \text{Table 3. DOLS estimates of the cointegrating coefficient} \]

| H0: \( \beta = 0 \) | 0.6120 |
| H0: \( \beta = 1 \) | 0.6120 |

\[ t\text{-ratio} = 3.4369 \]

\[ t\text{-ratio} = -2.1789 \]

\[ \text{Note: p-values are in parentheses.} \]

\[ \text{aThe DF test statistics are analogous to the parametric Dickey–Fuller test for nonstationary time series. The } DF_{rho} \text{ and } DF_t \text{ statistics assume strict exogeneity of the regressors with respect to errors and no autocorrelation. } DF_{rho}^* \text{ and } DF_t^* \text{ statistics are based on endogenous regressors. These tests depend on consistent estimates of the long-run variance–covariance matrix to correct for nuisance parameters once the limiting distribution has been found. The ADF test is analogous to the augmented Dickey–Fuller test for nonstationary time series.} \]

\[ \text{b} PC_1 \text{ and } PC_2 \text{ are the nonparametric Phillips–Perron tests.} \]

\[ 2 \text{ The Monte Carlo simulations presented in Kao and Chiang (2001) have shown that the DOLS estimator outperforms both the OLS and fully modified least squares estimators for both homogeneous and heterogeneous panels.} \]

\[ 3 \text{ On the other hand, the coverage of the fiscal data also excludes contingent fiscal liabilities that, if realized, will put strain on the public finances.} \]

\[ 4 \text{ See, for example, International Monetary Fund (2005a, b).} \]

\[ \text{III. Conclusions} \]

Panel cointegration techniques indicate that government revenue and expenditure in five Asian economies for the period 1974 to 2001 were nonstationary and cointegrated series. However, the cointegration coefficient was substantially less than unity, indicating ‘weak’ sustainability and the desirability for policy measures to put the public finances on a more sustainable basis. This would seem to justify the steps taken in the early part of this decade by many of these countries to strengthen their fiscal positions.

\[ \text{Table 2. Panel cointegration tests, 1974–2001} \]

<table>
<thead>
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<tbody>
<tr>
<td>( DF_{rho} )</td>
<td>( DF_t )</td>
</tr>
<tr>
<td>-4.8697 (0.0000)</td>
<td>-3.1675 (0.0000)</td>
</tr>
</tbody>
</table>

\[ \text{Note: p-values are in parenthesis.} \]

\[ a \text{The } DF \text{ test statistics are analogous to the parametric Dickey–Fuller test for nonstationary time series. The } DF_{rho} \text{ and } DF_t \text{ statistics assume strict exogeneity of the regressors with respect to errors and no autocorrelation. } DF_{rho}^* \text{ and } DF_t^* \text{ statistics are based on endogenous regressors. These tests depend on consistent estimates of the long-run variance–covariance matrix to correct for nuisance parameters once the limiting distribution has been found. The ADF test is analogous to the augmented Dickey–Fuller test for nonstationary time series.} \]

\[ b \text{ } PC_1 \text{ and } PC_2 \text{ are the nonparametric Phillips–Perron tests.} \]
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References


International Monetary Fund (2005b) *Regional Economic Outlook – Middle East and Central Asia*, September, IMF, Washington, DC.


