Tax-Spend, Spend-Tax, or Fiscal Synchronization: A Panel Analysis of the Chinese Provincial Real Data

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In this paper we tested whether the hypothesis of tax-spend, spend-tax, or fiscal synchronization applies to the 31 Chinese provinces using cross-sectional and time series data covering 1999 to 2005. The interaction between government revenues and government expenditures is tested with the newly developed panel unit root tests and heterogeneous panel cointegration tests. The results show that both revenues and expenditures are non-stationary but have a significant long-run relationship. The results based on multivariate panel error-correction models show that there is no significant causality between revenues and expenditures in the short run. However, in the long-run, a bi-directional causality exists between revenues and expenditures, thus supporting the fiscal synchronization hypothesis for 31 Chinese provinces over this sample period.

Keywords: tax-spend, spend-tax, fiscal synchronization, panel cointegration

JEL classification: C22, C23, H72

1 Introduction

Due to concerns over the growing budget deficits, numerous studies have been devoted to testing the “Tax-and-Spend”, “Spend-and-Tax”, and “Fiscal Synchronization” hypotheses. The “tax-spend” hypothesis suggests that changes in

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government revenues lead to changes in government expenditures and advocates that the government should attempt to re-balance the national budget by raising taxes. On the other hand, the “spend-tax” hypothesis suggests that changes in government expenditures lead to changes in government revenues and explains that additional government expenditures should be controlled or restricted in order to re-balance the budget. Finally, the “fiscal synchronization” hypothesis suggests that because expenditure and revenue decisions are often made simultaneously, both expenditures and revenues push the budget towards equilibrium.

Determining which hypothesis best characterizes an economy is more than an intellectual exercise because it can potential contribute towards the discovery of a solution to the problem of growing budget deficits. The amount of existing literature dedicated to the study expenditure – revenue relationships indicates the seriousness of research in this field. A series of country-specific studies are as follows: Anderson et al. (1986), Von Furstenberg et al. (1986), Miller and Russek (1990), and Baghestani and McNown (1994) for the US; Hasan and Lincoln (1997) for the UK; Payne (1997) for Canada; Darrat (1998) for Turkey; Li (2001) and Chang and Ho (2002b) for China; and Chang and Ho (2002a) for Taiwan. In addition, Ram (1988a, 1988b), Baffes and Shah (1994), Chang et al. (2002), and Reddick and Hassan (2003) conducted multi-country studies. Generally speaking, empirical evidence in testing the validity of these hypotheses has led to inconclusive results.

While previous studies focus predominately on industrial and developing countries, this study attempts to contribute to this line of research by using the newly developed panel unit root tests, panel cointegration tests, and panel-based error correction model to test the “Tax-Spend, Spend-Tax, or Fiscal Synchronization” hypothesis for the Chinese provinces.1

Before 1979, China’s fiscal system was characterized by centralized revenue collection and centralized fiscal transfers, meaning that most taxes and profits were collected by local governments, remitted to the central government, and then partially transferred back to the local governments based on their centrally-approved expenditure needs. Beginning in 1979, China began a market-oriented economic

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1The province level administrated by the People's Republic of China consisted of 22 provinces, 5 autonomous regions, 4 municipalities, and 2 special administrative regions, Honk Kong and Macau. In this study we focus on the provinces, autonomous regions, and municipalities directly under the central government, and denote them as “provinces” for simplicity.
reform. In particular, the tax reform was an important part of this economic reform, and it aimed at providing state enterprise production incentives, cutting off fiscal dependence of state enterprises on government, equalizing tax burdens among enterprises, and promoting fair competition. According to Lin (2000), the tax reforms occurred over five stages. In the first stage beginning in 1979, the central government experimented with allowing state enterprises to keep a partial portion of their profits. In the second stage, after the success of the experiments in stage 1, the government pursued further fiscal system reforms in 1983, adopting a “substituting taxes for profits” approach. In the third stage during December 1986, the contract responsibility system (CRS) was introduced on the basis of the “substituting taxes for profits” reform (for additional details about CRS see Lin (2000)). In the fourth stage, the government launched a tax plus profit system in 1989 to address the problem of declining government revenues. Finally, in the fifth stage, the government established a new tax system known as the “tax-sharing” system in 1994. Several other significant changes in the tax systems took place during this final stage (for more details about these tax reforms see Lin (2000)).

Overall, the current fiscal system in China emphasizes the sharing of tax revenues. Under the proposed system, expenditures of various levels of governments are based on their individually managed and balanced budgets. In addition, while the central government budget is approved by the National People’s Congress (NPC), local government budgets are approved by the People’s Congresses at the local levels. Central and local budgets are then divided into current items, capital, and/or construction items. Expenditure items include expenditures for social development, welfare, national defense, armed policies, and administration. Meanwhile, tax revenues make up 96% of total revenues.

While China’s per capita GDP has grown significantly since 1979, the nation’s budget deficit has also increased. By 1989, the Chinese government’s revenues fell below the necessary level required to balance the growing amount of government expenditures. The trend of growing budget deficits was observed in almost every Chinese province. The worsening economic situation can largely be explained by the following two reasons. First of all, provinces need extremely large amounts of infrastructure and public investments in order to operate and continue to develop, leading to rapid growth in government expenditures. Second, in most of the
provinces, government revenues have fallen beneath sufficient levels as taxable sources disappear, making tax collection increasingly difficult. Therefore, this paper examines whether there exists a force that causes government budgets to move from a state of deficit or deviation from the equilibrium to a long-run equilibrium for the Chinese provinces.

The data set used in this study consists of annual time series data on real government revenues and expenditures of 31 Chinese provinces over the period of 1999-2005. To begin our study, first we apply the panel unit root tests conducted by Levin, Lin and Chu (LLC) (2002), Im, Pesaran and Shin (IPS) (2003), and Hadri (2000) to examine the time series properties of the real government revenue and expenditure variables. Tests reveal that all the time series contain a unit root, indicating that all the real variables are non-stationary. The Pedroni (1997, 1999) heterogeneous panel cointegration test is subsequently used to examine whether a long-run equilibrium relationship exists among these two variables. The cointegration test suggests that these two variables are cointegrated. Finally, results based on multivariate panel error-correction models show that no significant causality exists between revenues and expenditures in the short run. However, in the long-run, a bi-directional causality exists between revenues and expenditures, thus supporting the fiscal synchronization hypothesis for Chinese Provinces.

This paper is organized as follows: Section II presents the data, Section III describes our methodology. Section IV discusses the empirical findings, and Section V presents our conclusions.

2 Data and Methodology

2.1 Data

In this empirical paper we use annual data on real government revenues and expenditures for 31 Chinese provinces over the 1999 to 2005 period (deflated by Provincial Retail Price Indices deflator, 1994 = 100). All the data used in this paper are taken from the China Statistical Yearbook (2000-2006), published by the National Bureau of Statistics of China.² The Retail Price Indices measure the degree

²Please refer to http://www.stats.gov.cn/tjsj/ndsj/.
and trends of changes in the retail prices between urban and suburban areas. Changes in retail prices have a direct impact on household expenditures and government revenues, and significantly influence the purchasing powers of the nation. In addition, change in retail prices disrupts equilibrium not only in market supply and demand, but also in national consumption and saving. Therefore, we use the Retail Price Indices as a deflator to fully reflect the real revenues and expenditures in each province of China.

2.2 Methodology

Ever since Nelson and Plosser (1982) published their seminal work, various studies have been devoted to investigating the potential non-stationarity of important macroeconomic variables. Researchers have been especially interested in the time-series properties of real output levels. As pointed out by Nelson and Plosser, the modeling of real output levels as either a trend stationary or difference stationary process has important implications for macroeconomic policy making, modeling, testing, and forecasting. Studies on this issue are of concern to not only empirical researchers, but also policymakers.

While numerous studies support a unit root in real output levels, critics claim that such conclusions may be attributed to the lower power of the employed conventional unit root tests. More recently, it has been reported that conventional unit root tests not only fail to consider information across regions, thereby leading to less efficient estimations, but also have lower power when compared to near-unit-root but stationary alternatives. It is not surprising that these factors have cast considerable doubt on many of the earlier findings that have been based on a unit root in real output levels. A tangible approach for increasing the testing power of the unit roots is to incorporate panel data. Recent literature suggests that the panel-based unit root tests, including the LLC (2002) test, the IPS (2003) test, and the Hadri (2000) test, have higher power than the traditional unit root tests based on time series. To account for heterogeneous panels, the Pedroni test (1997, 1999) is employed in this study. Finally, the panel vector error correction model is used to describe both long run relationships and short run dynamic adjustments between real government revenue and expenditure variables of the 31 Chinese provinces over the
period of 1999 to 2005.

2.2.1 Levin, Lin and Chu (2002) Panel Unit Root Test

LLC found that the panel approach substantially increases power in finite samples when compared with the single-equation ADF test. Based on the ADF specification, LLC proposed a panel-based version of equation (1) that restricts $\hat{\beta}_i$ by keeping it identical across cross-sectional regions as follows:

$$\Delta X_{it} = \alpha_i + \beta X_{it-1} + \sum_{j \neq i} \theta_j \Delta X_{jt-1} + \varepsilon_{it},$$

(1)

where $\Delta$ is the first difference operator, $X_{it}$ is the real provincial revenues and expenditures, $\varepsilon_{it}$ is a white noise disturbance with a variance of $\sigma^2$, $t = 1, 2, \ldots, T$ indexes time periods, and $i = 1, 2, \ldots, N$ indexes cross-sectional regions. LLC tested the null hypothesis for the existence of a unit root (i.e. the series is nonstationary) with $\beta_1 = \beta_2 = \cdots = \beta = 0$ against the one-side alternative of having no unit root with $\beta_i = \beta_2 = \cdots = \beta < 0$, based on the following test statistic:

$$t_{\beta} = \frac{\hat{\beta} - \beta}{se(\hat{\beta})},$$

where $\hat{\beta}$ is the OLS estimate of $\beta$ in equation (1), and $se(\hat{\beta})$ is its standard error. It is worth noting that the LLC test requires a specification of the number of lags used in each cross-section ADF regression, and that one must specify the exogenous variables used in the testing equations.

2.2.2 Im, Pesaran and Shin (2003) Test

IPS relaxed the assumption of identical first-order autoregressive coefficients of the LLC test and developed a panel-based unit root test that allows $\beta$ to be differed across regions under the alternative hypothesis. Meanwhile, IPS tested the null hypothesis of unit root with $\beta_1 = \beta_2 = \cdots = \beta = 0$ against the alternative of no unit root with $\beta_i < 0$, for some $i$. The IPS test is based on the mean group approach and they use the average of the $t_{\beta}$ statistics from equation (1) to perform the following
standardized t-bar statistic:

\[ Z = \frac{\sqrt{N} [I - \bar{E}]}{\sqrt{\text{Var}(I)}} \rightarrow N(0, 1), \]

(2)

where \( \bar{I} = (1/N) \sum_{i=1}^{N} i \), \( E(\bar{I}) \) and \( \text{Var}(\bar{I}) \) are the mean and variance for each \( i \) statistic, respectively. IPS has shown that \( Z \) has an asymptotic standard normal distribution, \( N(0, 1) \). Like the LLC test, the IPS test requires specification of the number of lags and the specification of the deterministic component for each cross section ADF equation. Based on the results of Monte Carlo Experiments, IPS (2003) demonstrated that the IPS panel unit root test is more powerful than the LLC panel unit root test.

2.2.3 Hadri (2000) Panel Unit Root Test

The Hadri (2000) panel unit root test is similar to the KPSS unit test, and has a null hypothesis of having no unit root in any of the series (i.e. the series is stationary) in the panel. Like the KPSS test, the Hadri test is based on the residuals from the individual OLS regressions on a single constant, or on a constant and a trend. Given that both constant term and trend are included, the following equation is estimated:

\[ y_s = \alpha_s + \beta_1 t + \epsilon_s, \]

where \( \epsilon_s \) is the residual from the regression, and the LM statistic is given as follows,

\[ LM_1 = \frac{1}{N} \left( \sum_{i=1}^{N} \left( \sum S \left( (t^2 / T^2) \right) / \tilde{\epsilon}_s \right) \right). \]

Allowing for heteroskedasticity across cross sections, we have an alternative LM statistic,

\[ LM_2 = \frac{1}{N} \left( \sum_{i=1}^{N} \left( \sum S \left( (t^2 / T^2) \right) / f_m \right) \right), \]

where, \( S(t) = \sum_{i=1}^{n} \tilde{\epsilon}_s \) is the cumulative sum of the regression residual and \( \tilde{\epsilon}_s = \sum_{i=1}^{n} f_m / N \) is the average of the individual estimators of the residual spectrum.
at a frequency of zero. Hadri (2000) shows that under mild assumptions,

\[ Z = \sqrt{N} \left( \frac{LM - \psi}{\xi} \right) \rightarrow N(0, 1), \]

where, \( \psi = 1/6 \) and \( \xi = 1/45 \) if the model includes only constants (\( \lambda \) is set to 0 for all \( i \)), and \( \psi = 1/15 \) and \( \xi = 11/6300 \), otherwise. The Hadri panel unit root test only requires the specification of the form of the OLS regressions: whether to include only individual specific constant terms, or to include both constant and trend terms. The results will be two \( Z \)-statistic values, one based on \( LM_1 \) with the associated homoskedasticity assumption, and the other based on \( LM_2 \), which is heteroskedasticity consistent.

### 2.2.4 Pedroni (1997, 1999) Heterogenous Panel Cointegration Tests

If two or more variables are cointegrated, then the time paths of one series must be influenced by the time paths of the other, to the extent that they cannot depart from each other for a long period of time. The Johansen procedure is commonly used to test for the existence of cointegration between variables. However, the power of the traditional Johansen cointegration approach is severely inhibited when applied to a small sample size. To address this weakness, we need to combine information from time series and cross section data before employing panel cointegration tests. Pedroni (1997, 1999) developed a number of panel cointegration test statistics based on the residuals of the Engle and Granger (1987) study. They allow for the heterogeneity among individual members of the panel, include heterogeneity in both the long run cointegrating vectors and in the short dynamics, and does not impose any exogeneity requirements on the regressors in the cointegrating regressions. By so doing, Pedroni derived seven panel cointegration statistics for varying intercepts and varying slopes to test the null hypothesis where there is no cointegration among heterogenous panels. The first category among four statistics collectively known as the pooled panel cointegration statistics, is defined as the “within-dimension-based statistics” and includes a variance ratio statistic (panel \( \nu \)-statistic), a non-parametric Phillips and Perron type rho statistic (panel \( \rho \)-statistic), a non-parametric Phillips and Perron type t-statistic (panel PP-statistic) and an
Augmented Dickey-Fuller type t-statistic (panel ADF-statistic). The second category of three panel cointegration statistics is defined as the “between-dimension-based statistics” and is based on a group mean approach. The set includes a Phillips and Perron type rho-statistic (group \( \rho \)-statistic), a Phillips and Perron type t-statistic (group PP-statistic), and an Augmented Dickey-Fuller type t-statistic (group ADF-statistic). Following Pedroni (1997, 1999), the heterogeneous pooled panel cointegration test statistics are calculated as follows:

\[
\text{Panel } \nu\text{-statistic } Z_i = \left( \sum_{t=1}^{T} \sum_{i=1}^{N} \hat{L}_{it}^2 \hat{\epsilon}_{i,t}^2 \right)^{-1},
\]

\[
\text{Panel } \rho\text{-statistic } Z_\rho = \left( \sum_{t=1}^{T} \sum_{i=1}^{N} \hat{L}_{it}^2 \hat{\epsilon}_{i,t}^2 \right)^{-1/2} \sum_{i=1}^{N} \sum_{t=1}^{T} \hat{L}_{it} \hat{\epsilon}_{i,t+1} \Delta \hat{e}_t - \hat{\lambda},
\]

\[
\text{Panel PP-statistic } Z_{PP} = \left( \sum_{t=1}^{T} \sum_{i=1}^{N} \hat{L}_{it}^2 \hat{\epsilon}_{i,t}^2 \right)^{-1/2} \sum_{i=1}^{N} \sum_{t=1}^{T} \hat{L}_{it} \hat{\epsilon}_{i,t+1} \Delta \hat{e}_t - \hat{\lambda},
\]

\[
\text{Panel ADF-statistic } Z_A = \left( \sum_{t=1}^{T} \sum_{i=1}^{N} \hat{L}_{it}^2 \hat{\epsilon}_{i,t}^2 \right)^{-1/2} \sum_{i=1}^{N} \sum_{t=1}^{T} \hat{L}_{it} \hat{\epsilon}_{i,t+1} \Delta \hat{e}_t.
\]

The heterogeneous group mean panel cointegration test statistics are as follows:

\[
\text{Group } \rho\text{-statistic } Z_\rho = \sum_{i=1}^{N} \left( \sum_{t=1}^{T} \hat{\epsilon}_{i,t}^2 \right)^{-1/2} \sum_{i=1}^{N} \hat{\epsilon}_{i,t+1} \Delta \hat{e}_t - \hat{\lambda},
\]

\[
\text{Group PP-statistic } Z_{PP} = \sum_{i=1}^{N} \left( \sum_{t=1}^{T} \hat{\epsilon}_{i,t}^2 \right)^{-1/2} \sum_{i=1}^{N} \hat{\epsilon}_{i,t+1} \Delta \hat{e}_t - \hat{\lambda},
\]

\[
\text{Group ADF-statistic } Z_A = \sum_{i=1}^{N} \left( \sum_{t=1}^{T} \hat{\epsilon}_{i,t}^2 \right)^{-1} \sum_{i=1}^{N} \hat{\epsilon}_{i,t+1} \Delta \hat{e}_t.
\]

where \( \hat{\sigma}^2 \) is the pooled long-run variance for non-parametric model given as \( 1/N \sum_{i=1}^{N} \hat{\sigma}_i^2 \), and \( \hat{\lambda} = 1/2(\hat{\sigma}^2 - \hat{\sigma}_i^2) \), where \( \hat{\lambda} \) is used to adjust for autocorrelation in panel parametric model, \( \hat{\sigma}^2 \) and \( \hat{\sigma}_i^2 \) are the long-run and contemporaneous variances for individual \( i \), and \( \hat{\sigma}_i^2 \) is obtained from individual ADF-test of \( e_{i,t} = \rho e_{i,t-1} + v_{i,t} \). \( \hat{\sigma}_i^2 \) is the individual contemporaneous variance from the parametric model, \( \hat{\epsilon}_{i,t} \) is the estimated residual from the parametric cointegration, while \( \hat{\epsilon}_{i,t}^* \) is the estimated residual from parametric model. \( \hat{L}_{it} \) is the estimated long-run covariance matrix for \( \Delta \hat{e}_{i,t} \) and \( L_{it} \) is the \( t \)th component of low triangular Cholesky decomposition of matrix \( \Omega \) for \( \Delta \hat{e}_{i,t} \) with the appropriate lag length determined by the New-West method.
The asymptotic distributions of these statistics are derived in Pedroni (1997). Pedroni argues that for cases with longer time spans (such as, $T > 100$), the sample size distortion tends to be insignificant while retaining a very high testing power across all seven statistics. For shorter panels, however, the alternative statistics might yield conflicting evidence. Pedroni shows that in terms of testing power, the group-ADF statistic has the best general performance, followed by the panel-ADF, while the panel-variance and the group-rho statistics have the worst performance.

2.2.5 Panel Vector Error Correction Model

As Granger (1986) pointed out, if cointegration exists between variables, then there is at least one causal relationship among them in one direction, suggesting that the Granger causality tests can be used to examine the nature of such relationships. A vector error correction model (VECM) is a restricted VAR designed to be used with nonstationary series that are known to be cointegrated. The VECM has cointegration built into its specification so that it restricts the long run behavior of the endogenous variables in order to converge to their cointegrating relationships while allowing for short run adjustment dynamics. The cointegration term is known as the correction term since deviations from the long-run equilibrium are corrected gradually through a series of partial short run adjustments.

If real government revenues and expenditures in this study are cointegrated, then a long run relationship exists between them. We can use the vector error correction model to characterize both long run equilibrium relationships and short run dynamic adjustment processes between government revenues and government expenditures. To construct an error correction model for the 31 provinces, a vector error correction model with heterogeneous panels is established as follows:

\[
\Delta GR_{it} = \theta_{it} + \sum_{k=1}^{k} \theta_{k} \Delta GR_{i,t-k} + \sum_{k=1}^{k} \theta_{ik} \Delta GE_{i,t-k} + \lambda_i \varepsilon_{i,t} + \eta_{i,t},
\]

\[
\Delta GE_{it} = \phi_{it} + \sum_{k=1}^{k} \phi_{ik} \Delta GR_{i,t-k} + \sum_{k=1}^{k} \phi_{ik} \Delta GE_{i,t-k} + \lambda_i \varepsilon_{i,t} + \nu_{i,t},
\]

where, $\Delta$ is the first difference operator, $k$ is the lag length, $GR_{it}$ and $GE_{it}$ are the real provincial revenues and expenditures respectively. Both $\varepsilon_{i,t}$ and $\varepsilon_{i,t-1}$ are the error correction terms. $\varepsilon_{i,t} = GR_{i,t} - \beta_i GE_{i,t-1} - \alpha_i$ and $\varepsilon_{i,t-1} = GE_{i,t-1} - \beta_i GR_{i,t-1} - \alpha_i$. Parameter $\lambda_i$ is the speed of adjustment to long-run
equilibrium. $\eta_t$ and $\nu_t$ are the statistical noises. This model can be estimated using instrumental variables to deal with the correction between the error term and the lagged dependent variables.

3 Empirical Results

To test for the existence of a unit root in a panel data setting, we used tests based on those conducted by LLC (2002), IPS (2003), and Hadri (2000). Results of the panel unit root tests are reported in Table 1. The LLC and IPS panel data unit root tests, with and without a deterministic trend component, support the hypothesis of nonstationarity of government revenues and government expenditures. Based on the Hadri unit root tests, with and without a time trend, we reject the null hypothesis of stationarity of government revenues and government expenditures respectively. All of the three panel unit root tests indicate the presence of unit roots in the time series of provincial revenues and expenditures, therefore the provincial budget data are nonstationary in levels for all 31 provinces in China.

<table>
<thead>
<tr>
<th>Test</th>
<th>GR</th>
<th>GE</th>
</tr>
</thead>
<tbody>
<tr>
<td>LLC</td>
<td></td>
<td></td>
</tr>
<tr>
<td>without time trend</td>
<td>19.7424</td>
<td>9.81363</td>
</tr>
<tr>
<td>with time trend</td>
<td>22.9379</td>
<td>–8.35408</td>
</tr>
<tr>
<td>IPS</td>
<td></td>
<td></td>
</tr>
<tr>
<td>without time trend</td>
<td>10.5385</td>
<td>7.63247</td>
</tr>
<tr>
<td>with time trend</td>
<td>6.51035</td>
<td>–5.60548</td>
</tr>
<tr>
<td>Hadri</td>
<td></td>
<td></td>
</tr>
<tr>
<td>without time trend</td>
<td>10.5677***</td>
<td>10.4986***</td>
</tr>
<tr>
<td>with time trend</td>
<td>13.6017***</td>
<td>16.1727***</td>
</tr>
</tbody>
</table>

Notes: *** donates significance at 1% level.

Panel cointegration tests are used in order to draw sharp inferences since time spans of economic time series are typically short. The estimated Pedroni’s test statistics are given in Table 2. The results presented in Table 2 show that the null
hypothesis of having no cointegration is rejected by the seven test statistics at the 5% significance level when testing for cointegration between provincial revenues and provincial expenditures. Thus, there is a long-run relationship between provincial revenues and provincial expenditures. These results suggest that the provincial expenditures are helpful in explaining the behavior of provincial government revenues in China in the long run, and vice versa.

### Table 2: Panel cointegration tests

<table>
<thead>
<tr>
<th>Test</th>
<th>Dependent Variable</th>
</tr>
</thead>
<tbody>
<tr>
<td>Test GR</td>
<td>GE</td>
</tr>
<tr>
<td>Panel variance</td>
<td>4.430672***</td>
</tr>
<tr>
<td>Panel $\rho$</td>
<td>$-2.552867**$</td>
</tr>
<tr>
<td>Panel ADF</td>
<td>$-4.834225***$</td>
</tr>
<tr>
<td>Panel PP</td>
<td>$-5.620879***$</td>
</tr>
<tr>
<td>Group $\rho$</td>
<td>2.622935**</td>
</tr>
<tr>
<td>Group PP</td>
<td>$-5.993482***$</td>
</tr>
<tr>
<td>Group ADF</td>
<td>$-11.25982***$</td>
</tr>
</tbody>
</table>

Notes: ***, **, and * donate significance at 1%, 5%, and 10% levels respectively.

The empirical results of the panel error correction model are reported in Table 3. According to the Table 3, lagged provincial expenditures have a negative impact on current provincial revenues. In other words, an increase in lagged provincial expenditures will cause a decrease in current provincial revenues. However, the effect is not significant. On the other hand, lagged provincial revenues have a positive impact on the current provincial expenditures. An increase in lagged provincial revenues will cause an increase in current provincial expenditures but the effect of provincial revenues on provincial expenditure is also insignificant. Even though provincial expenditures have a negative effect on provincial revenues and provincial revenues have a positive effect on provincial expenditures, both effects are insignificant. Therefore, there is no strong evidence to support the claim that short run causality exists between provincial revenues and provincial expenditures. Table 3 also indicates that there exists a significant cointegrating relationship between provincial revenues and provincial expenditures because the estimates of $\lambda_1$ and $\lambda_2$ are significant. Since $\lambda_1 < 0$, this implies that if provincial
Table 3: Panel vector error correction model

<table>
<thead>
<tr>
<th>Independent Variable</th>
<th>ΔGR (-1)</th>
<th>ΔGE (-1)</th>
<th>ΔGR (-2)</th>
<th>ΔGE (-2)</th>
<th>Error correction term</th>
</tr>
</thead>
<tbody>
<tr>
<td>ΔGR (-1)</td>
<td>0.560</td>
<td>0.897</td>
<td>0.866</td>
<td>1.036</td>
<td>-0.195**</td>
</tr>
<tr>
<td></td>
<td>(0.166)</td>
<td>(0.151)</td>
<td>(0.177)</td>
<td>(0.161)</td>
<td>(0.036)</td>
</tr>
<tr>
<td>ΔGE (-1)</td>
<td>-0.217</td>
<td>-0.515</td>
<td>-0.351</td>
<td>-0.668</td>
<td>-0.341**</td>
</tr>
<tr>
<td></td>
<td>(0.138)</td>
<td>(0.125)</td>
<td>(0.132)</td>
<td>(0.120)</td>
<td>(0.033)</td>
</tr>
</tbody>
</table>

Notes: p-values in parenthesis. ** denotes significance at 5% level.

expenditures in the previous period have overshot the equilibrium (i.e. $e_{i,t-1} < 0$), then the error correction term induces a positive changes in provincial revenues (GR) back towards equilibrium. Similarly, $\lambda_i < 0$ implies that if provincial expenditures in the previous period have overshot the equilibrium (i.e. $e_{i,t-1} > 0$), then the error correction term works to push provincial expenditures (GE) back towards the equilibrium. Both revenues and expenditures adjust in response to deviations between the two variables and will approach long-run equilibrium eventually. This observation suggests that long-run bi-causality exists between government revenues and government expenditures for China’s 31 provinces. In addition, these empirical results also support the fiscal synchronization hypothesis which states that tax and spending decisions are made simultaneously by the fiscal authority in China’s provinces over the observed sample period. The results match our assumptions and claims about China’s fiscal system mentioned earlier in this paper. The major implication that we draw from our results is that in order to attack the problem of continuously increasing budget deficits, the provincial government of China should be cautious, as pointed out by Manage and Marlow (1986), about simply raising revenue, cutting expenditures, or simply changing both revenues and expenditures without taking into consideration that the interdependence between the two may be ambiguous in their impacts on fiscal situation in China. While our empirical findings are consistent with similar studies on China by Li (2001), and Chang and Ho
(2002b), they are inconsistent with those found in Chang and Ho (2002a)’s study on Taiwan and Chang et al. (2002)’s study based on a panel of 7 industrial countries and 3 newly industrialized countries. Such studies suggest that unidirectional Granger causality exists between government revenues and government expenditures for most studied countries. This difference may well reflect the different fiscal system used in China and those of other countries (see, Ma (1997), Luo and Golembiewski (1996), Lin (2000), and Li (2001)).

4 Conclusions

In this paper we tested the hypothesis of tax-spend, spend-tax, and fiscal synchronization for 31 Chinese provinces using cross-sectional and time series data covering 1999 to 2005.

Our study improves upon research in this area in several respects. First, existing studies focus mostly on developing and developed countries instead of looking at China and her centrally planned economy. In addition, we use a newly developed panel unit root test, the LLC test (2002), the IPS test (2003), and the Hadri test (2000), and the heterogeneous panel cointegration tests from Pedroni (1997, 1999) to analyze and test the interactions between government revenues and expenditures. The results show that while both provincial revenues and provincial expenditures in China are non-stationary, they share a significant long-run relationship. On the other hand, empirical evidence from multivariate panel error-correction models show that there is no significant causality between provincial government revenues and expenditure in the short run. Nevertheless, in the long run, bi-directional causality exists between provincial government revenues and expenditures, suggesting that both provincial government revenues and expenditures help push the budget towards equilibrium. This finding support the fiscal synchronization hypothesis for the 31 Chinese provinces analyzed over this sample period. The results of fiscal synchronization in this study demonstrate the impact of coordination between institutions on budgetary outcomes. This being that both sides of the budget must be coordinated and separate institutions are responsible for this at the provincial level.
References


