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Does the expansionary government spending crowd out the private consumption? Cointegration analysis in panel data

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Abstract

That whether expansionary government spending crowds out private consumption is examined by evaluating the intra-temporal elasticity of substitution between them. Using annual data (1981–2000) of 23 OECD countries, a linear deterministic cointegration relation between private consumption, government spending and their relative price is supported. We have two findings: first, the panel estimators plausibly compute the parameter estimates in general. Second, when cross-sectional correlation is considered by using a SUR estimator, the statistical significance of panel cointegration is improved. Thirdly, the *intra*-temporal elasticity of substitution indicates that government and private consumption are found to be complements, which shows that expansionary government spending does not crowd out private consumption.

JEL classification: C22; E21; E62

Keywords: Panel cointegration; Two-step estimator; SUR; Intratemporal substitution

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1. Introduction

This paper investigates whether an increase in government spending, or any other exogenous increase in spending, has a greater *substitutability* effect on the nominal level of income through price increases, real income increases, or both. This observation makes expansionary fiscal policy attractive to those who believe in government intervention to improve the economy. In a permanent income model, unlike the multiplier analysis emphasizing the short-run dynamics, this paper focuses on investigating whether the government spending and the private consumption are complements or substitutes.

Financing an increase in government spending by increasing taxes or by selling bonds to the public causes crowding-out forces to reduce the magnitude of the impact. Although financing by printing money may avert crowding-out effect temporarily, it will cause inflation and raise the interest rate in the future. Whether the increase in national income is due to the increase in the money supply or the increase in government spending is still a debatable issue. Bailey (1971) indicated that there might be a degree of substitutability between government spending and private consumption. Barro (1981) incorporated it into a general model to examine the direct effect of government purchases of goods and services on consumption utility. Kormendi (1983) and Aschauer (1985) applied the permanent-income approach and find a significant degree of substitutability between private consumption and government spending for the United States. Ahmed (1986) estimated the effects of UK government consumption in an intertemporal substitution model and found that government expenditures tend to crowd out private consumption. Recently, Aiyagari, Rao, Christiano, & Eichenbaum (1992) and Baxter and King (1993) explored the effect of government spending shocks on various economic aggregates in a one-sector neoclassical growth model with constant returns to scale and variable labor supply. They find that increases in government spending significantly led to a decline in private consumption.

Some studies show different results. Karras (1994) examines the change of private consumption in response to increases in government spending across a number of countries and finds that public and private consumption are better described as complementary rather as substitutes. The strength of this complementary relationship is shown to be negatively affected by the government size. These findings are robust across all specifications. In other words, in the aggregate, they are best described as complementary goods in the sense that an increase in government spending tends to raise the marginal utility of private consumption. In terms of a neoclassical model with increasing returns to scale and monopolistic competition, Devereux, Allen, & Beverly (1996) examined the impact of government spending shocks and found that an increase in government consumption generates an endogenous rise in aggregate productivity. The increase in productivity raises the real wage sufficiently that there is a substitution away from leisure and into consumption. Thus, an increase in government expenditures leads to an increase in private consumption. These findings imply that private consumption cannot be responsible for any crowding-out effects that government spending might have on aggregate demand. On the contrary, private consumption is probably “crowded-in”.

This paper attempts to revisit this issue by contributing to the literature in two important aspects: first, the effectiveness of fiscal expansion is evaluated as the intratemporal elasticity of substitution between government spending and private consumption, so an *inelastic* intratemporal (elasticity of) substitution of government spending for private consumption implies a smaller coefficient than it is anticipated. We employ Ogaki's (1992) model to estimate the intratemporal elasticity of substitution between government spending and private consumption. In this model, cointegration is applied to exploit the long run restriction imposed by the intraperiod first order condition. This method can directly evaluate the intratemporal elasticity of substitution from the data that are shown to be robust to a number of economic factors. Second, we extend the line of research to panel data. The fundamental advantage of a panel data set over a cross section is that it allows the researcher far greater flexibility in modeling differences in behavior across individuals (see Baltagi, 1995; Hsiao, 1996).

There are two explanations justifying the empirical application of the representative agent model to panel data. First, sometimes it is plausible to believe that different countries should have some common features in our sampling period. Common features could be the number of common trends or that the cointegrating relations should lie in the same space. That is, it result implies that the long-run equilibrium should hold for each cross-sectional country, but it is feasible to allow the cross-sections to depend on all the equilibrium of the cross-sections. Thus, second, if the common-feature is not plausible, then the presence of cointegrating relation will not hold for the panel data. One of the original motivations to test for stationarity in panel data is due to the lack power of conventional univariate unit root tests against persistent alternatives, typically for sample sizes that occur in practice. This has also been proven quite satisfactory in improving the power of unit root tests. The additional cross-sectional dimension in the panel leads to better power properties of the panel tests as compared to the lower power of the standard individual-specific unit root test against near unit root alternatives for small samples.

Recent developments in non-stationary panel data have sparked a large body of literature. Quah (1994) is one of those who pioneer the research, who proposes the tests that exploit information from cross-sectional dimensions in inferring non-stationarity from panel data. Pedroni (1996, 1997) proposed a fully modified estimator for heterogeneous panels, and derive asymptotic distributions for residual based tests of cointegration for both homogeneous and heterogeneous panels. Kao (1999), Hadri (2000), Levin, Lin, and Chu (2002) and Im, Pesaran, and Shin (2003) constituted further important contributions along the line. Kao (1999) first examined the behavior of spurious panel regression. He has provided an asymptotic theory for the behavior of LSDV estimator in a model with $I(1)$ variables and show that the OLS estimator is consistent for its true value, but the t-statistic diverges so that the inferences about the regression coefficient are wrong with a probability that goes to 1. Hence, in the presence of panel unit root, we have to account for non-stationary panel data and test panel cointegration. To this end, Kao (1999) proposed residual-based tests for the null hypothesis of no panel cointegration. Kao and Chiang (2000) proposed several algorithms to conduct regression for non-stationary panel data. Recently, Breitung (2002) further contributed to this area by proposing a two-step estimator to estimate the long-run cointegrating vectors.

This paper is organized as follows. The next section develops the theoretical model. Section 3 describes the data and some basic time series properties. Section 4 analyzes the cointegrated panel. Section 5 concludes.

2. The theoretical model

Following Ogaki (1992) and Amano and Wirjanto (1997), let C_t and G_t denote the real per capita private expenditures and government consumption at period t, respectively. Thus, a representative country who maximizes the expected lifetime consumption utility function is expressed below

$$U = E_t \left[\sum_{t=0}^{\infty} \beta^t U(C_t, G_t) \right] \tag{1}$$

Eq. (1) is subject to a lifetime budget constraint in a complete market at period 0. E_t is the expectation operator based on information of period t, $0 \leq \beta \leq 1$ is a discount factor. Consider the addilog function:

$$V(C_t, G_t) = \left(\frac{C_t^{1-\alpha} - 1}{1-\alpha} \right) \Lambda_{C_t} + S \left(\frac{G_t^{1-\nu} - 1}{1-\nu} \right) \Lambda_{G_t} \tag{2}$$

where α and ν are curvature parameters with $\alpha, \nu \geq 0$, S is a scaling factor and Λ_C and Λ_G represent random preference shocks associated with private and public consumption, respectively. Obviously, Eq. (2) is a constant relative risk aversion (CRRA) function, and $1/\alpha$, and $1/\nu$ are interpreted as the inter-temporal elasticity of substitution for private and government consumption, respectively. As allowing for these shocks, we avoid Garber and King's (1983) assertion that the presence of random preference shocks can often yield misleading results. Hence, the representative consumer maximizes the intraperiod utility of Eq. (2) subject to the intratemporal budget constraint below

$$P_c(t)C_t + P_g(t)G_t = M(t)$$

where $P_c(t)$ and $P_g(t)$ are the prices of private consumption and public consumption at time t. and $M(t)$ is the total consumption expenditure at time t. To exploit the empirical implications of the model, Ogaki (1992) assumes an intratemporal utility function given by

$$U(C_t, G_t) = f_t[V(C_t, G_t)], \quad f_t' > 0 \tag{3}$$

where f_t is an arbitrary monotonic transformation function. In addition, for the empirical analysis, we also assume that the sequences of random preference shock are stationary or $I(0)$ processes. An intratemporal (or static) first-order necessary condition of the above-mentioned problem states that the relative purchase price of government to private expenditures is equated to the marginal rate of substitution based on the purchase of the two types of goods, namely,

$$P_t = \frac{G_t^{-\nu} S \Lambda_{G_t}}{C_t^{-\alpha} \Lambda_{C_t}} \tag{4}$$

Eq. (4) can be written as

$$G_t S^{1/\nu} P^{1/\nu} C_t^{-\alpha/\nu} = \left(\frac{\Lambda_{G_t}}{\Lambda_{C_t}} \right)^{1/\nu} \tag{5}$$

Let the lower case letters denote variables in logarithmic form and taking logarithms on both sides of Eq. (5) yields

$$c_t = \mu + \frac{1}{\alpha} p_t + \frac{\nu}{\alpha} g_t + \varepsilon_t \tag{6}$$

where $\mu = 1/\alpha \ln S_t$ and ε_t $(=1/\alpha \ln(\Lambda_{G,t}/\Lambda_{C,t}))$ is a stationary process of preference shocks with zero mean. Due to the methodology of representative individual country, we assume that all cross-sectional units have similar preference parameters; hence, the panel form of Eq. (6) is hence specified below

$$c_{it} = \mu_i + \frac{1}{\alpha} p_{it} + \frac{\nu}{\alpha} g_{it} + \varepsilon_{it} \tag{7}$$

Eq. (7) has three important implications: first, the coefficients associated with the second term are interpreted as the intratemporal elasticity of substitution. That is, ν/α measures the contemporaneous causality between the two variables. The government spending affects the private consumption by ν/α. Hence, the effectiveness of a fiscal expansion is determined by the magnitude of parameter estimate. Second, this equation also allows us to estimate the inter-temporal elasticity of substitution simultaneously; that is, $1/\alpha$. According to Amano and Wirjanto (1997), there are three testable implications for (7),

(a) If $1/\alpha > \nu/\alpha$, then C_t and G_t are Edgeworth-Pareto complements, under which, the *intertemporal* elasticity of substitution is larger than the *intratemporal* elasticity of substitution.
(b) If $1/\alpha < \nu/\alpha$, then C_t and G_t are Edgeworth-Pareto substitutes, under which, the *intratemporal* elasticity of substitution is larger than the *intertemporal* elasticity of substitution.
(c) If $1/\alpha = \nu/\alpha$, then C_t and G_t are Edgeworth independent, or unrelated.

3. Data and some time series properties

Data of 23 OECD countries is derived from the *AREMOS/OECD*, a database compiled by the Ministry of Education (Taiwan, Republic of China). To make our empirical results has meaningful implications for current policies, the sampling periods should not be too long; hence, the sample period extends from 1981 to 2000. The private consumption includes consumer spending on goods and services. The government consumption consists in government spending on goods and services. The real variables are calculated by the current price and exchange rate of 1980, and expressed by millions of U.S. dollar. All per capita variables are obtained by dividing the aggregate measure by total population. Table 1 offers a list of the 23 economies, and country annually averages over 1981–2000 for the three variables. Figs. 1–3 present the time plots of them. We visually examine the stationarity of each series. Since 23 countries are too many to be plotted in a single graph, we put eight countries in one figure. It looks that both per capita private consumption and government consumption have an increasing trend over time, however, we do not know whether it is a stochastic trend (unit root) or deterministic trend. Next, we conduct formal statistical tests for panel unit root.

Table 1
Sample means, 1981–1997

Countries	P_t	C_t	G_t
Australia	0.042552	0.009292	0.002611
Austria	−0.00848	0.011117	0.003814
Belgium	0.026912	0.011811	0.002731
Canada	−0.01027	0.010547	0.004189
Denmark	−0.00723	0.012284	0.006328
Finland	−0.06778	0.012172	0.005005
France	0.008308	0.011548	0.003616
Germany	−0.01158	0.011482	0.004083
Greece	−0.03423	0.006261	0.001237
Iceland	0.038554	0.011348	0.003564
Ireland	0.009382	0.007194	0.001895
Italy	−0.07584	0.010811	0.003142
Japan	−0.0186	0.013024	0.002122
Luxembourg	−0.02299	0.015460	0.003426
Norway	0.031329	0.011086	0.002734
The Netherlands	0.00597	0.006906	0.002857
New Zealand	−0.00768	0.012673	0.003340
Portugal	0.002907	0.004125	0.001021
Spain	0.009737	0.007275	0.001829
Sweden	−0.04728	0.012306	0.006702
Switzerland	−0.02363	0.018535	0.004374
UK	−0.03662	0.009735	0.003349
USA	0.009969	0.013924	0.003513

Note: AREMOS/OECD Data Set, Ministry of Education, Taiwan, ROC (1999).

Table 2 reports the unit root tests for individual countries. We conduct two types of tests with different null hypothesis: the ADF (Dickey & Fuller, 1981; Said & Dickey, 1984) tests the unit-root-null, and the KPSS (Kwiatkowski, Phillips, Schmidt, & Shin, 1992) tests the no-unit-root-null. The reason why we employ two unit root test with different null hypothesis is that: if the observed series are clearly trending, the low power will primarily be due to the strong multicollinearity in the ADF regression between deterministic trend and the lagged variable. An alternative way to handle this potential multicollinearity is to switch the null and alternative hypothesis and consider the stationarity test, such as the KPSS test. Under the null of KPSS the series are specified to be trend stationary, and under the alternative the series are difference stationary. A comparison of the results obtained from the ADF and KPSS tests can give some insights into the stationarity properties of the series. If both the ADF and the KPSS tests fail to reject the null hypotheses (or, if both reject), we have mixed results and can only conclude that the data are not informative enough. But, on the other hand, if the ADF tests fail to reject the null and the KPSS tests reject, we have more confidence in the results that the series under consideration are in fact non-stationarity. The same procedure applies to tests in panel data.

Reading across the rows of the table is the country-by-country results for 23 OECD countries. Clearly, the presence of unit root in each country is unanimously confirmed by ADF, and the KPSS rejects the stationarity null most. To continue the study, we first test

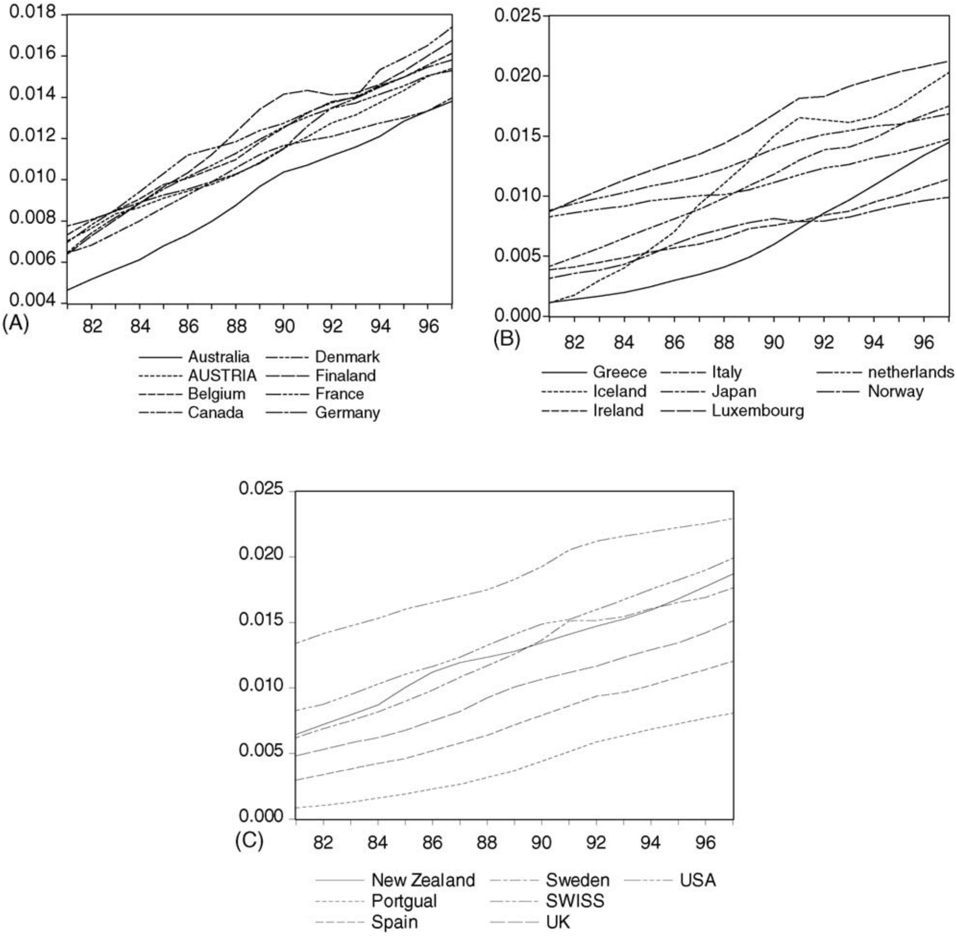

Fig. 1. (A) Per capita private consumption. (B) Per capita private consumption. (C) Per capita private consumption.

for panel unit root. Three panel unit root tests mentioned previously are applied: IPS (Im et al., 2003), HT (Harris & Tzavalis, 1999), and the KPSS-based LM (Hadri, 2000).

Since IPS is directly based upon conventional ADF, we do not introduce its technical contents here. The LM statistic tests the null that the panel series is stationary. First, consider the following dynamic data generating processes (Hadri, 2000):

$$\text{Level}: y_{it} = r_{it} + \varepsilon_{it} \tag{8}$$

$$\text{Trend}: y_{it} = r_{it} + \beta_i t + \varepsilon_{it} \tag{9}$$

Here r_{it} is a random walk: $r_{it} = r_{it-1} + \theta \cdot u_{it}$, and both u_{it} and r_{it} are generated from $N(0,1)$. The stationary null hypothesis is simply expressed by H_o: $\sigma_u^2 = 0$. Specifically, the null hypothesis is specified below $H_o : \rho = \frac{\sigma_u^2}{\sigma_\varepsilon^2} = 0, \;\; \text{against} \;\; H_a : \rho > 0$

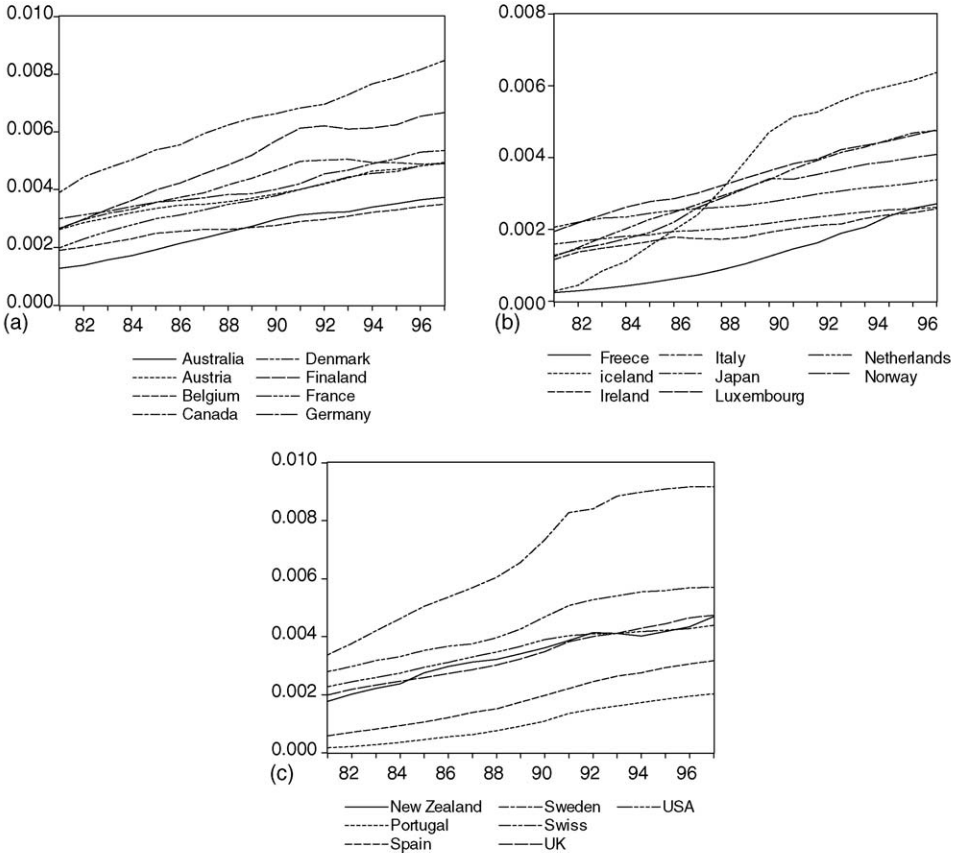

Fig. 2. (A) Per capital government consumption. (B) Per capital government consumption. (C) Per capital government consumption.

The LM statistic used to test the null hypothesis is defined below

$$\mathrm{LM} = \frac{\sum_{i=1}^{N}\sum_{t=1}^{T} S_{it}^2}{N \cdot T^2 \cdot \varpi^2}, \quad S_{it} = \sum_{j=1}^{t} \varepsilon_{ij} \tag{10}$$

where ϖ^2 is a consistent estimator of σ_u^2 under the null.

The HT statistic tests the null that there is a panel unit root. HT assumes the time dimension is fixed and is derived from the following models

$$\text{Level} : y_{it} = \alpha_i + \varphi y_{it-1} + \nu_{it} \tag{11}$$

$$\text{Trend} : y_{it} = \alpha_i + \beta_i t + \varphi y_{it-1} + \nu_{it} \tag{12}$$

$$H_{\mathrm{o}} : j = 1$$

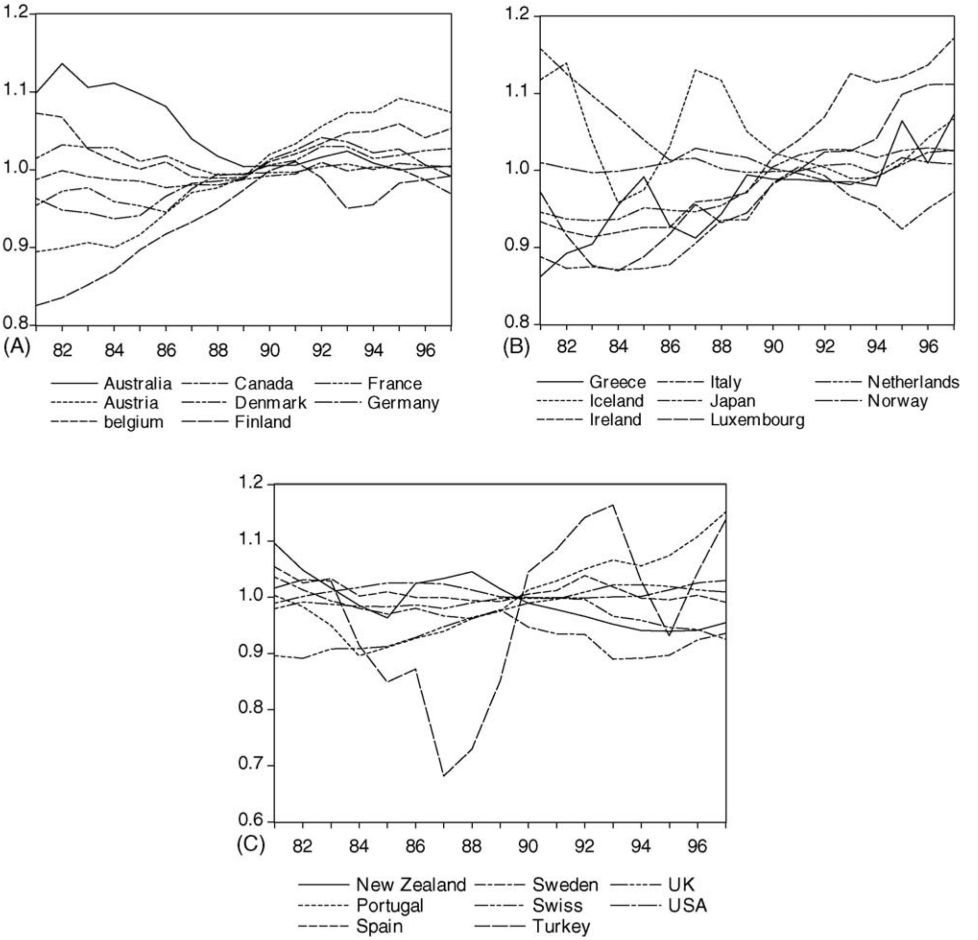

Fig. 3. (A) Relative price indices. (B) Relative price indices. (C) Relative price indices.

Under the null hypothesis, the LS estimator of φ satisfies

$$\hat{\varphi} - 1 = \left[\sum_{i=1}^{N} y'_{i,t-1} Q_T y_{i,t-1}\right]^{-1} \cdot \left[\sum_{i=1}^{N} y'_{i,t-1} Q_T v_{it}\right] \tag{13}$$

where Q_T is the appropriate $T \times T$ idempotent transformation matrix.

Table 3 summarizes the estimation outcomes. To avoid small-sample bias, we calculate the critical values of the panel-based test by using Monte-Carlo simulations calibrated for our sample size. According to the critical values presented in Table 5, the presence of panel unit root is confirmed by three tests.

Table 2
Unit root tests of individual country

Country	p_t			c_t			g_t		
	ADF(1)	KPSS		ADF(1)	KPSS		ADF(1)	KPSS	
		level	trend		level	trend		level	trend
Australia	−1.120	0.541	0.093	−0.979	0.546	0.093	−1.460	0.194	0.179
Austria	−0.940	0.542	0.082	−0.924	0.463	0.092	−0.674	0.189	0.142
Belgium	−0.440	0.531	0.084	−1.293	0.449	0.093	−0.021	0.233	0.151
Canada	−0.790	0.523	0.117	−0.711	0.461	0.118	−0.816	0.181	0.139
Denmark	0.370	0.547	0.088	0.120	0.477	0.096	−1.507	0.163	0.141
Finland	−0.980	0.472	0.121	−1.041	0.375	0.126	−1.089	0.142	0.141
France	−1.080	0.532	0.121	−1.571	0.414	0.127	−1.130	0.134	0.132
Germany	−1.260	0.531	0.122	−1.622	0.404	0.111	−1.060	0.138	0.130
Greece	−0.590	0.541	0.124	−0.620	0.464	0.095	−0.865	0.186	0.142
Iceland	−0.970	0.511	0.117	−1.511	0.390	0.118	−1.252	0.308	0.167
Ireland	1.990	0.534	0.158	1.722	0.416	0.163	−1.014	0.360	0.154
Italy	−1.630	0.524	0.126	−1.863	0.395	0.127	−0.671	0.359	0.147
Japan	−1.710	0.526	0.130	−1.711	0.416	0.132	−0.487	0.311	0.170
Luxembourg	−0.930	0.534	0.097	−1.552	0.522	0.095	−0.603	0.396	0.157
The Netherlands	0.240	0.539	0.126	0.142	0.534	0.114	−0.585	0.313	0.155
New Zealand	0.980	0.515	0.118	0.541	0.529	0.086	−0.881	0.186	0.158
Norway	2.060	0.546	0.114	−0.512	0.527	0.085	−0.924	0.320	0.168
Portugal	−2.010	0.518	0.110	−1.393	0.512	0.093	−0.500	0.326	0.169
Spain	−1.060	0.527	0.118	−1.361	0.508	0.094	−1.228	0.297	0.183
Sweden	−0.910	0.514	0.125	−1.283	0.440	0.119	−0.688	0.262	0.173
Switzerland	−1.560	0.496	0.127	−1.081	0.514	0.133	−1.098	0.236	0.179
UK	−1.270	0.528	0.112	−1.563	0.520	0.120	−1.647	0.195	0.183
USA	0.310	0.543	0.097	−0.352	0.544	0.118	−0.790	0.264	0.167

Note: p in ADF(p) denotes the number of lags. ADF are calculated with trend and intercept. The MacKinnon (1991) critical values for ADF at 1, 5, and 10% significance level are −4.044, −3.451, and −3.151. For KPSS tests, the approximate asymptotic critical values at 10, 5, and 1% significance level for the level model are, respectively, 0.347, 0.463, and 0.739; and for the trend model, they are 0.119, 0.146, and 0.216, respectively.

Table 3
Tests for panel unit-root

Test statistic	p_{it}	c_{it}	g_{it}
LM			
Level	0.748	0.817	0.791
Trend	0.469	0.587	0.512
IPS			
Level	−1.13	−0.97	−1.05
Trend	−1.06	−1.14	−1.12
HT			
Level	1.19	1.06	1.15
Trend	1.21	1.12	1.07

Note: critical values are listed in Table 5 below.

4. Estimation of non-stationary panel

4.1. Kao and Chiang (2000)

Under the assumption of homogeneous long-run covariance across cross-sectional units, Kao and Chiang (2000) propose three models which produce asymptotically efficient estimators for the long-run cointegrating vector $\boldsymbol{\beta}$ in the presence of cross-sectional dependence: the bias-corrected OLS, the fully-modified OLS, and the dynamic OLS. Since the bias-corrected OLS has poor performance, we do not use it here. To illustrate, let X and Y denote the independent and dependent variables, respectively,

$$\hat{\beta}_{\mathrm{FM}} = \left[\sum_{i=1}^{N}\left(\sum_{t=1}^{T}(x_{it}-\bar{x}_i)\hat{y}_{it}^{+} - T\hat{\Delta}_{\varepsilon u}^{+}\right)\right]\left[\sum_{i=1}^{N}\sum_{t=1}^{T}(x_{it}-\bar{x}_i)(x_{it}-\bar{x}_i)'\right]^{-1} \tag{14}$$

$$y_{it} = \alpha_i + x'_{it}\hat{\beta}_{\mathrm{DOLS}} + \sum_{j=-q}^{q} c_{ij}\Delta x_{it+j} + v_{it} \tag{15}$$

Obviously, testing for panel cointegration of the model is equivalent to testing the stationarity of the residuals.

4.2. Breitung's (2002) two-step estimator

Breitung (2002) proposes a two-step estimator, which is based on a cointegrated VAR(p) model. To explain this procedure, consider a VECM model

$$\Delta y_{\mathrm{it}} = \alpha_i\beta' y_{i,t-1} + \varepsilon_{it}, \quad i = 1, 2, \ldots, N;\ \ t = 1, 2, \ldots, T \tag{16}$$

where ε_{it} is a k-dimensional white noise error vector with $E(\varepsilon_{it}) = 0$ and positive definite covariance matrix $\Sigma_i = E(\varepsilon_{it}\cdot\varepsilon_{it-1})$. The long-run parameter β can be obtained conditional on some consistent initial estimator of α_i and Σ_i. To derive the two-step estimator, Breitung (2002) transforms (16) into a VECM below

$$\gamma'_i\Delta y_{\mathrm{it}} = \gamma'_i\alpha_i\beta' y_{i,t-1} + \gamma'_i\varepsilon_{it} \tag{17}$$

$$z_{it} = \beta' y_{i,t-1} + \nu_{it} \tag{18}$$

where $z_{it} = (\gamma'_i\alpha_i)^{-1}\gamma'_i\Delta y_{it}$, $\nu_{it} = (\gamma'_i\alpha_i)^{-1}\gamma'_i\varepsilon_{it}$ and γ is a $k\times r$ matrix with $rk(\gamma'_i\alpha_i) = r$. It follows that $\Sigma_\nu - (\alpha'_i\Sigma^{-1}\alpha_i)^{-1}$ is positive semi-definite and, therefore, the optimal choice of the transformation is $\gamma'_i = \alpha'_i\Sigma_i^{-1}$. The resulting estimator is asymptotically equivalent to the Gaussian ML estimator.

At first stage, the individual equation of (17) is estimated by VAR and conducts a normalization procedure to obtain α_i and Σ_i. The restriction that the cointegrating vectors are the same for all cross-sectional units is ignored, but this does not affect the asymptotic properties of the estimator. For the asymptotic properties of the two-step estimator, it is only required that the parameters are estimated consistently as T approaches infinity.

At the second estimation stage, the system is transformed into (18) such that the cointegration matrix β can be estimated by least-squares of the pooled regression. Table 4 presents

Table 4
Parameter estimates of panel cointegration

Variables	FMOLS	DOLS (1,1)	Two-step estimator
Cross-sectional correlation is absent			
p_t	1.240 (40.94)	1.72(15.53)	−1.249(1.14)
g_t	0.789 (14.25)	0.751(31.71)	0.663 (5.10)
Panel cointegration test			
LM			
Level	0.398	0.384	0.117
Trend	0.317	0.464	0.027
Variables	**FM-SUR**	**Dynamic SUR**	**Two-step SUR**
Cross-sectional correlation is present: SUR estimator			
p_t	1.199 (2.14)	1.211 (2.06)	−1.109 (8.94)
g_t	0.972 (16.43)	0.874 (14.42)	0.712 (9.77)
Panel cointegration test			
LM			
Level	0.119	0.209	0.077
Trend	0.092	0.064	0.048

Note: numbers in the parentheses are the student t-statistics. The LM statistics for panel cointegration are significant at 1% significance level.

the estimation results. We first examine the LM statistics for panel cointegration. As we note previously, the residual-based tests for cointegration is subject to the hypothesis under the null. If the null for the residuals is of the ADF type, then under the null, the model is not cointegrated, and the simulation of critical values is not plausible. Hence, we test the null that the model is cointegrated and the KPSS-based LM statistic is employed. According to the critical values listed in Table 5, except two-step estimator, the FMOLS and DOLS either strongly rejects the null or weakly accepts it.

The parameter estimates for the FMOLS and DOLS estimators are, respectively, (1.240, 0.789) and (1.721, 0.751). The theoretical prediction implied by (7) indicates that the both parameters are positive; hence, both models yield plausible inter-temporal elasticity of substitution ($1/\alpha$) and intra-temporal elasticity of substitution (ν/α). Under which, the *intertemporal* elasticity of substitution is greater than the *intratemporal* elasticity of substitution; therefore, we can interpret it as that the government spending and private spending are Edgeworth-Pareto complements.

Table 5
Simulated critical values for panel data

Significance level (%)	LM		IPS		HT	
	Level	Trend	Level	Trend	Level	Trend
1	0.511	0.198	−1.389	−1.382	2.217	2.212
5	0.329	0.123	−1.211	−1.228	1.875	1.616
10	0.211	0.097	−1.081	−1.084	1.453	1.417

Note: 20,000 replications with $N = 23$ and $T = 17$. IPS and HT are almost the same for both specifications.

The essence of the panel time series in the literature is that we take averages across time and countries. Taking averages over time may not be serious; however, averages over cross-sectional units can be very sensitive to outliers. To check this, we report the country-by-country estimates at the tables of the Appendix A, showing that the parameters ν/α are very close. However, the parameter estimate $1/\alpha$ varies widely across the countries and some of them do not have the sign predicted by the theory. Therefore, although there are some weak points, most of the absolute value of $1/\alpha$ are larger than ν/α and the majority of 23 OECD countries have pausible results; hence, it may still justify the inference regarding their relationship in a framework of panel data.

Hence, we keep on examining the outcomes estimated by Breitung's (2002) two-step estimator, and we find that Breitung's two-step estimator poorly explores the data. The parameter estimates are (−1.249, 0.663), the first parameter estimate (−1.249) does not have correct sign. This problem may be a result of the presence of outliers. In addition, cross-sectional correlation can be important. One of the crucial issues in panel data model is the possible cross-sectional dependence on parameter estimates. It is known that unaccounted cross-sectional dependence may cause severe distortion in statistical inferences. As suggested by Breitung (2002), this can be easily solved by applying a SUR estimator to the second stage estimation, because the second stage estimation is based on a pooled LS regression. For FMOLS and DOLS, we use Moon's (1999) FM-SUR and the Dynamic SUR of Mark, Ogaki, and Sul (2003). FM-SUR can be calculated by employing a SUR estimator to fully-modified data, so is the Dynamic SUR.

The bottom panel of Table 4 presents the estimation results and we have important findings. First, the parameter sings are plausibly obtained by both FM and DOLS. The parameter estimates for the FMSUR and DSUR estimators are, respectively, (1.199, 0.972) and (1.211, 0.874). Both models do yield plausible $1/\alpha$ and ν/α. Because $\nu/\alpha < 1/\alpha$, under which, the *intertemporal* elasticity of substitution is greater than the *intratemporal* elasticity of substitution; therefore, we can interpret it as that, for the countries in the panel, the government spending and private spending are Edgeworth-Pareto complements. In a nutshell, the expansionary fiscal policy may be effective. However, the Breitung's two-step estimator obtains (−1.109, 0.712) the sign of $1/\alpha$ does not conform to theoretical prediction.

Besides, unlike those in the top panel, the LM statistics for FM-SUR and Dynamic SUR accepts the null of panel cointegration at 5% significance level.

5. Conclusions

For the countries in the panel, the government spending and private spending are Edgeworth-Pareto complements. Therefore, our empirical results justify the Keynesian plea for expansionary fiscal policy.

Three points are summarized as follows: firstly, cross-sectional correlation is important to FMOLS and DOLS, both models estimate plausible parameter estimate of inter-temporal elasticity of substitution and the absence of cointegration; when cross-sectional correlation is controlled by SUR, panel cointegration is significantly supported by FMOLS and DOLS. Second, the according to country-by-country estimates (shown in the Appendix A), the inter-temporal elasticity of substitution varies across 23 countries, but the intra-temporal elasticity of substitution does not differ much across 23 economies. Finally, we find Breitung's (2002) two-step estimator is not robust in exploring the theoretical prediction underlying the data.

Despite the success of our analysis, the empirical framework of this paper has two weak points for research in the future: first, according to Graham (1993), it is necessary to examine the stability of the parameter estimates of the cointegrating vector. Unfortunately, the stability test for non-stationary panel data regression has not been developed satisfactorily so far, and conventional stability tests for single-equation cointegration model cannot be directly extended to panel cointegration model. Second, as the tables in the appendix indicate, although the intra-temporal elasticity of substitution (ν/α) of 23 OECD member countries lies in a homogeneous numerical interval, the inter-temporal elasticity of substitution ($1/\alpha$) is rather diverse; thus, exploring this problem is important. They are left as research in the future.

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Appendix A

See Tables A.1 and A.2.

Table A.1
Country-by-country estimation results, FMOLS

Country	$1/\alpha$		ν/α		AIC
	Estimate	S.D.	Estimate	S.D.	
Australia	−1.692	0.286	0.774	0.003	−4.403
Austria	1.180	0.032	0.808	0.000	−3.628
Belgium	1.103	0.649	0.760	0.004	−4.300
Canada	−0.946	2.123	0.824	0.006	−4.484
Denmark	2.173	0.399	0.866	0.002	−4.294
Finland	0.767	0.070	0.823	0.001	−3.977
France	−3.567	0.521	0.787	0.001	−3.600
Germany	2.552	0.181	0.809	0.001	−3.765
Greece	4.453	0.929	0.741	0.008	−3.907
Iceland	−0.458	0.582	0.785	0.006	−3.562
Ireland	1.682	0.158	0.794	0.002	−4.425
Italy	2.629	0.455	0.750	0.007	−4.156
Japan	2.413	0.126	0.700	0.001	−4.282
Luxembourg	0.817	0.103	0.735	0.002	−4.197
The Netherlands	−0.885	0.536	0.761	0.004	−4.027
New Zealand	2.347	0.711	0.850	0.002	−4.358

Table A.1 (*Continued*)

Country	$1/\alpha$		ν/α		AIC
	Estimate	S.D.	Estimate	S.D.	
Norway	−2.043	0.399	0.765	0.003	−3.864
Portugal	1.065	0.046	0.800	0.000	−3.770
Spain	−0.920	0.665	0.775	0.002	−4.139
Sweden	−0.933	0.226	0.888	0.003	−4.325
Switzerland	−0.075	0.113	0.733	0.001	−4.421
UK	2.770	0.197	0.800	0.002	−3.749
USA	4.317	3.853	0.747	0.011	−3.521

Table A.2
Country-by-country estimation results, DOLS

Country	$1/\alpha$		ν/α		AIC
	Estimate	S.D.	Estimate	S.D.	
Australia	0.657	0.583	1.093	0.116	−3.781
Austria	0.714	0.361	0.988	0.165	−4.242
Belgium	1.696	0.436	1.053	0.088	−3.743
Canada	−2.939	1.438	1.185	0.069	−3.898
Denmark	−1.064	0.284	1.302	0.049	−4.000
Finland	0.407	0.386	0.991	0.114	−4.093
France	−1.833	0.354	0.912	0.030	−3.943
Germany	0.956	0.132	1.156	0.029	−3.738
Greece	−0.615	0.327	1.015	0.024	−4.206
Iceland	1.225	0.919	0.915	0.128	−4.022
Ireland	1.074	0.684	1.072	0.344	−3.789
Italy	0.224	0.206	1.221	0.077	−4.314
Japan	0.665	0.535	1.211	0.169	−3.703
Luxembourg	0.105	0.319	1.074	0.105	−3.939
The Netherlands	0.419	0.108	1.370	0.026	−4.275
New Zealand	2.377	0.705	0.994	0.026	−4.238
Norway	0.156	0.402	1.057	0.070	−4.319
Portugal	0.308	0.095	0.881	0.012	−3.778
Spain	0.142	0.370	0.802	0.027	−4.500
Sweden	2.097	0.689	1.240	0.067	−4.142
Switzerland	0.178	0.110	0.710	0.006	−4.051
UK	−0.185	1.559	1.417	0.326	−4.168
USA	1.142	0.927	1.325	0.079	−4.047

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