

Assessing the Twin Deficits Hypothesis in Selected OECD Countries: An Empirical Investigation

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Abstract: *The improvement of budget deficits and current account deficits has renewed the debate over the twin deficits hypothesis. This paper examines the validity of the twin deficits hypothesis in nine members of the Organizations for Economic Co-operation and Development (OECD) over the period 1990-2007. We use recently developed panel cointegration tests with structural breaks to assess the relationship between budget deficits and current account deficits. The empirical findings in this paper are as follows: (1) There is a long-run relationship between budget deficits and current account deficits for these countries; (2) The empirical result is consistent with the Keynesian proposition.*

Keywords: Twin deficits, panel unit root, structural breaks, panel cointegration

JEL Classification: C23, E6, F4

1. Introduction

In recent years, many scholars have studied the relationship between budget deficits and current account deficits in both developed and developing countries. In the past decade, the twin deficits phenomenon has reemerged in many countries. The importance of twin deficits stems from the “harmful” effects of deficits on an economy and the threats to macroeconomic stability. The two deficits are depending on the underlying tax system, trade patterns and barriers, the exchange rate and a complex host of internal and international forces that shape a country’s economic status in the global setting. The twin deficit hypothesis first emerged during the Reagan period of the 1980s in U.S., the budget deficit rose from 0.5% of full employment GDP in 1981 to 4.2% in 1985. The trade deficit rose over this period from 0.1% of GDP to 2.6% of GDP. The recession in 2001 mirrored the 1980s experience both in terms of the budget deficit and the trade deficit (Labonte, 2003). During the middle of the 2000s, some argued that a larger fiscal deficit, through its effect on national savings, would rekindle the current account deficit (Bartolini and Lahiri, 2006). During the 1980s and 2000s, the main cause of the U.S. current account deficit was the federal government budget deficit. The lack of savings in the government and private sectors must be complemented by foreign savings, resulting in capital inflow (Ito, 2009).

The United States is not a unique country from the point of view of twin deficits. Although the existing literature mostly focuses on the U.S., similar problems exist in both developed countries, such as Germany, England and Sweden, and in developing countries, such as Malaysia, Philippines and Turkey. These country currencies are dependent on the U.S. dollar. With the appreciation of the dollar, the increase in their budget deficits during the

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early 1990s led to current account imbalances, which resulted in an economic crisis (Baharumshah et al., 2006). The outbreak of the current global financial crisis has not wiped away the twin deficits issue. Indeed, many international economists blame it on the global crisis (Ito, 2009).

The aim of this article is to explore the validity of the twin deficits hypothesis in nine OECD countries using panel cointegration tests with structural breaks from 1990 to 2007. This article differs from existing related studies in three ways. First, although some studies have used trade deficits in the twin deficits hypothesis, this paper includes the current account of the balance of payments. The current account deficit, more comprehensive than the trade deficit, is the sum of the balance of trade (the export of goods and services minus the import goods and services) and net current transfers (interests, dividends, and foreign aid). Second, although there is now significant evidence to support the twin deficits hypothesis (Piersanti, 2000; Normandin, 1999; Chinn and Ito, 2005), none of these previous studies is based on panel cointegration with structural breaks. Most studies use time series data (Holmes, 2009; Grier and Ye, 2009) or classical panel data in their modeling (Saleh et al., 2005; Baharumshah et al., 2006); there are not structural breaks in either the unit root tests or the cointegration analysis. Third, some researchers have focused their attention on the relationship between the two deficits in one country, such as the United States, Egypt, Korea or Greece. A study of one country certainly cannot lead to an empirical generalization. To fill these gaps in the literature and to capitalize on the attention paid to the impact of the current economic crisis on budget deficits and current account deficits, this article examines the relationship between budget deficits and current account deficits in nine OECD countries using panel cointegration with structural breaks tests developed by Banerjee and Carrion-i-Silvestre (2006).

The structure of this article is as follows. Section 2 reviews the existing theory and literature. Section 3 explains the details of the methodology and presents the empirical results. Section 4 concludes with a summary of our results.

2. Theory and Relevant Literature

2.1. Theory

The twin deficits hypothesis refers to a causal relationship, namely that an increase in a budget deficit will lead to an increase in the current account deficit. Budget deficits and current account deficits have important policy implications for a number of reasons. First, persistent large deficits cause indebtedness by borrowing internally and externally. Second, deficits impose a burden on future generations (Hakro, 2009). Third, these deficits are harmful to related countries and the world economy. These deficits may damage the foreign exchange markets and cause high real interest rates, crises in international financial markets and low savings rates (Daly and Siddiki, 2009; Barro, 1989). Additionally, the current account deficit is an important indicator of economic performance; it is closely related to the budget balance and private savings, which are key factors of economic growth (Aristovnik, 2008). As a result, both deficits imply a lower level of wealth for economies. For these reasons, the current account deficit and the budget deficit must be under control to sustain economic growth. It is often difficult, however, for countries to maintain a sufficient level of control (Kouassi et al., 2004).

Some economists look at budget deficits and current account deficits from different perspectives. For example, Hashemzadeh and Wilson (2006) suggest that a large current account deficit may harm employment and production in some sectors of the economy, but it may also create an inflow of foreign capital and increase employment and spending opportunities in other sectors (Hashemzadeh and Wilson, 2006). However, it is generally accepted that the twin deficits have harmful effects on an economy. Both developed and developing countries have suffered from large budget and current account deficits to varying degrees.

In the literature, the theoretical debates on the relationship between budget deficits and current account deficits are based on two major theoretical models: the Keynesian proposition and the Ricardian equivalence hypothesis (REH, hereafter).

2.1.1. The Keynesian Proposition

The Keynesian proposition presents a positive relationship between budget deficits and current account deficits. It states that persistent budget deficits cause current account deficits, which indicates unidirectional causality. That is, the Keynesian proposition supports the twin deficits hypothesis. The Keynesian proposition is also referred to as “the conventional wisdom” or “the traditional view” (Enders and Lee, 1990; Piersanti, 2000).

In the Keynesian proposition, budget deficits have important and harmful effects on an economy. The harmful effects include high interest rates, low savings and low rates of economic growth.

The Keynesian proposition is based on the Mundell-Fleming framework (Kouassi et al., 2004: 504). In the Mundell-Fleming framework and under a flexible exchange rate, an increase in budget deficits would result in upward pressure on real interest rates and inflation and increase aggregate demand because the economy would be operating below full employment capacity. When a budget deficit decreases, the government adds to the national savings supplied by households and businesses, and interest rates fall. On the contrary, the growing budget deficits represent a claim on those savings, and interest rates must rise for the market to remain in equilibrium (Labonte, 2003). Second, in an open economy, due to the increase in the world interest rates (Blanchard, 1985), high interest rates cause inflows of foreign capital (short-term debt), and the domestic currency will appreciate. Third, the stronger currency reduces net exports due to relatively cheaper import prices. Consequently, current account imbalances will arise. Persistent budget deficits will widen current account deficits under both fixed and flexible exchange rate regimes (Anoruo and Ramchander, 1998). In other words, debt-financed expansionary fiscal policy is directly related to current account deficits (Leachman and Francis, 2002).

In addition, capital inflows can be reversed. A sudden reversal in large capital inflows leads to large fluctuations in macroeconomic variables such as domestic savings, exchange rate, interest rate and current account deficits (Baharumshah et al., 2006).

According to Feldstein and Horioka (1980), savings and investment rates are highly correlated, causing budget and current account deficits. Feldstein and Horioka reflect high capital mobility. If national saving decreases, it should not “crowd out” domestic investment. Feldstein and Horioka’s puzzle suggests that a strong correlation between investments and savings in advanced economies.

They used below equation to assess the relation between savings and investment rates:

$$\left(\frac{I}{Y}\right)_i = \alpha + \beta \left(\frac{S}{Y}\right)_i \quad (1)$$

Where $\left(\frac{I}{Y}\right)_i$ is the ratio of gross domestic investment to gross domestic product and $\left(\frac{S}{Y}\right)_i$ is corresponding ratio of gross domestic saving to ratio of gross domestic. α , the absolute term of the formula. β , investment's sensitive to saving increase. β coefficient nearing 0 shows perfect international capital mobility, β , nearing 1 shows the lack of the international capital mobility (Misztal, 2011).

In an open economy, the theoretical framework of the twin deficits hypothesis may exist under the auspices of a national account identity:

$$Y = C + I + G + X - M \quad (2)$$

$$CAD = X - M - NT \quad (3)$$

The current account is related to savings and investment in an economy. The current account balance requires a surplus of private savings over investment. In an open economy, national savings (S) can be aggregated into the following form:

$$S = Y - C - G + CAD \quad (4)$$

$$\text{Where } Y - C - G = I \quad (5)$$

$$S = I + CAD \quad (6)$$

$$S = S^p + S^g \quad (7)$$

$$S^p = Y - T - C, \quad S^g = T - G \quad (8)$$

To see the relationship between a budget deficit and current account deficit, from the national accounting identity (6) we obtain:

$$CAD = (S^p - I) - (G - T) \quad (9)$$

where Y is the gross domestic product (GDP), C is private consumption, I is private investment, S^p is private saving, S^g is government saving, S is total national savings, G is government expenditure, T is taxes collected from households and firms by the government, X is the export of goods and services, M is the import of goods and services, NT is net current transfers from abroad and CAD is the current account deficit. The budget deficit is given by $(G - T)$. An increase in the government deficit will increase the current account deficit. $(S^p - I)$ is unchanged and represents the private savings deficit. An increase in temporary purchases will increase the budget deficit, which directly affects the current account (Vamvoukas, 1999).

In a closed economy, net exports ($X - M$) are zero and private domestic savings equals private domestic investment. In an open economy, such a relationship may not always exist (Saleh et al., 2005).

In equation (9), when the private savings deficit ($S^p - I$) remains stable and public savings declines, an increase in the budget deficit leads to an increase in the current account deficit. For a given level of private savings and investment, the government budget and current account move in the same direction and by the same amount. This relationship is known as the twin deficits hypothesis (Anoruo and Ramchander, 1998; Daly and Siddiki, 2009; Baharumshah et al., 2006).

When we focus on variable ($G - T$) in equation (9), a tax increase will reduce the budget deficit and will lead to a reduction in private sector expenditures, which reduces the current account deficit. Conversely, a tax cut or other fiscal expansion that is financed by the issuance of public debt reduces national savings by increasing private disposable income and, as a result, private consumption (Bartolini and Lahiri, 2006).

Many countries have been running large budget deficits as a consequence of expansive fiscal policies. In this theoretical framework, budget deficits have three potential causes:

1) When taxes remain stable, an increase in public expenditures causes a budget deficit. Additionally, the increase of public consumption expenditures will increase national income by the multiplier effect. Higher national income results in an appreciation of domestic currency and increased imports. Alkswani (2000) and Beetsma et al. (2008) find a negative relationship between budget deficits and public expenditures. In the European Union, an increase in public expenditures of 1% of GDP leads to a decrease in the trade balance of 0.5% of GDP and a decrease in the budget balance of 0.7% of GDP. Together, these results provide support for the twin deficits hypothesis.

2) National savings equals private savings plus public savings (see equation (7)). A decrease in public savings diminishes national savings, which determines savings capacity. This leads to disequilibrium between savings (S) and investment (I). As result, a decrease in savings' capacity to finance investment supports the budget deficits hypothesis. Offsetting domestic savings through foreign direct investment will increase external government debt. Increasing external debt leads to budget deficits. Although a budget deficit provides a surplus in the capital account of the balance of payments due to the flow of foreign direct investment, it leads to a deficit in the current account.

3) The final cause of budget deficits is a decrease in tax revenues. Also, the causes of both public expenditures and tax revenues may stem from unsatisfied management and enterprise development.

The main causes of budget deficits vary among countries. For example, in the 1980s, large budget deficits in the U.S. were created primarily by increased military expenditures, high interest rates and a decline in the savings rates (Alkswani, 2000).

2.1.2. The Ricardian Equivalence Hypothesis

The REH contests the Keynesian proposition. The Ricardian equivalence suggests that the current account deficit is independent of the budget deficit. The REH is supported by

Barro (1989), who rejects any link between the two deficits. Although budget deficits or government debt increase aggregate demand and wealth, the Ricardian equivalence suggests that budget deficits do not affect wealth. The debate over Ricardian equivalence held a prominent position in the economic literature until the end of the 1980s (Ricciuti, 2003).

The REH states that an expansionary fiscal policy has no effect on consumption and output. According to the REH, budget deficits arise from tax cuts, which reduce public revenues. A reduction in current taxes creates an increase neither in consumption nor in national savings. Under this hypothesis, it is assumed that public expenditures are constant. This hypothesis can be explained in two ways. First, this hypothesis assumes that people rationally believe that current tax cuts are temporary. Decreased taxes in the current period will inevitably be balanced by future tax increases. In other words, a decrease in current taxes must be matched by an increase in the present value of future taxes. Fiscal policy would affect aggregate demand if it altered the expected present value of taxes. However, the present value of taxes would not change if the present value of spending changed. People believe that they will have to pay more tax to compensate for budget deficits in future. Rational people will not change their private consumption. Therefore, budget deficits and taxation have equivalent effects on the economy. This situation is referred to as the "Ricardian equivalence theorem" (Reitschuler, 2008). The REH reveals that tax and debt have the same effect on private consumption. Proponents of the REH argue that government debt represents a future tax liability. The substitution of taxes for government debt does not create current account deficits (Enders and Lee, 1990). Therefore, an increase in disposable income, due to decreased taxes, does not lead to an increase in consumption.

Second, taxes cut do not affect national savings because the reduction in public savings is offset by an increase in private savings (Ricciuti, 2003). In addition, according to the REH, the shift between taxes and budget deficits does not affect real interest rates. This is contrary to the Keynesian view, which presents a positive effect (Barro, 1989). Decreased public savings creates budget deficits, but the decrease in public savings will be matched by an equal increase in private savings. The total savings level is not affected. As a result, budget deficits have no effect on current account deficits (Alkswani, 2000). In short, given a specific level of public and private spending, a tax increase will reduce budget deficits, but current account deficits will not be affected (Enders and Lee, 1990).

Barro (1989) has suggested five primary theoretical objections to the REH. First, people do not consider taxes that are levied in the future because people do not live forever. Second, private capital markets are imperfect. Third, future taxes and other incomes are uncertain. Fourth, all tax incomes are not lump sum; most taxes are determined by income, spending and wealth. The fifth and final reason is that the REH is based on full employment.

Échevin (2009) tested the Ricardian view via a survey questionnaire that aimed to assess the impact of the 2002 French tax cut on private consumption. He found that the proportion of "spender" consumers, who spent a large portion of their tax cut, was high relative to "Ricardian" consumers, who did not. Approximately 52.7% of the households stated that the tax reduction led to an increase in their consumption. Additionally, Échevin (2009) found evidence that the average marginal propensity to consume tax cuts (76.5%) is significantly greater than the average marginal propensity to consume temporary increases in earnings (42.4%). According to Échevin (2009), most of households answered that they consumed the tax cut, while another significant share said that they did not know how they spent it. Assuming that non-responses are equally divided between the various modes, and

that, for those who say they have both consumed and saved it (or used it for other purposes), half of the tax cut is actually consumed and another half is saved, then the consumption ratio of the tax cut (i.e. the proportion of the tax cut that has been consumed) amounts to 81.5 per cent. In contrast, lower-income households preferred to save a larger proportion of the tax cut (Échevin, 2009).

The REH may explain some issues in various countries. For example, in the U.S. economy, a decrease in savings during the last twenty years can be partially clarified by the Ricardian equivalence. The contractionary fiscal policy in some small European countries can also be explained by this equivalence (Ricciuti, 2003).

2.2. Relevant Literature

In economic literature, there have been extensive empirical studies that examine the relationship between budget deficits and current accounts using different models. Each study contributes its own explanation. Although a consensus in the literature has not yet emerged on the existence and direction of a causal relationship between budget deficits and current account deficits, the twin deficits debate has helped to expand our understanding of the macroeconomic consequences of large budget and trade deficits. The different results in the literature stem from variations in the empirical technique, data measurement and sample set (Rosenswieg and Tallman, 1993). Many empirical research efforts have focused on the twin deficits in developed countries, while few empirical studies have examined developing countries. For example, Anoruo and Ramchander (1998) study the twin deficits hypothesis for developing Southeast Asian countries. The study of Baharumshah et al. (2006) is based focused on four developing Asian countries. Daly and Siddiki investigate the impact of budget deficits and real interest rates on current account deficits for 23 developed OECD countries. According to their results, previous empirical research can be classified into five categories:

First, some studies find support for the REH. Barro (1989), Enders and Lee (1990) and Evans (1988) have rejected the existence of a relationship between budget deficits and current account deficits using U.S. data. They argue that the shift between tax and debt finance does not affect aggregate demand. For this reason, output, price levels, interest rates and exchange rates should not be affected by budget deficits.

Second, a number of studies provide evidence for the Keynesian proposition. That is, they found a relationship between budget deficits and current account deficits. Most of these studies are based on data from developed countries. Piersanti (2000), who used the Granger-Sims causality technique, found strong evidence that future budget deficits had a positive effect on trade deficits for 17 OECD countries over the period 1970-1997. The research performed by Normandin (1999) in Canada and the U.S. supports the twin deficits hypothesis. This hypothesis implies that a tax increase would directly decrease the budget deficit and indirectly reduce the current account deficit. This implies a decrease in the consumption of imported goods induced by the decrease in private after-tax incomes. Similarly, Rosenswieg and Tallman (1993) investigated whether fiscal policy plays an essential role in the adjustment of trade imbalances in the U.S., and they found a link between budget and trade deficits.

The empirical results of Baharumshah et al. support the twin deficits hypothesis for Thailand in the long term. They find a direct causal link between budget deficits and current account deficits and an indirect relationship between budget deficits and higher interest

rates. Higher interest rates lead to an appreciation of the country's currency and widen the current account deficit. This chain of causal relationships is valid in all countries (Thailand, Malaysia and Indonesia) included in the analysis except the Philippines (Baharumshah et al., 2006). Saleh et al. (2005), in their empirical analysis, found support for the Keynesian view in the long -run.

Blanchard (1985) suggests a positive correlation between budget deficits and current account deficits. Blanchard rejected the Ricardian argument that a decrease in taxes today does not influence private consumption and wealth. Under a finite time horizon, contrary to the infinite horizon assumed by Ricardo, taxes are partly shifted to a future generation. Therefore, someone currently alive will not have to pay the future increase in taxes (Blanchard, 1985). Doménech et al. (2000) test the REH for the sample of OECD countries. Their results do not support the Ricardian equivalence hypothesis. Private savings compensates for only a fraction of the budget deficit. The main reason for the decline in public savings is the budget deficit. The deterioration of public savings has not been balanced by an equivalent increase in private sector savings. Chinn and Ito (2005) find that the budget balance is an important determinant of the current account balance for industrial countries.

Although many studies of the twin deficits are based on short-run analysis, Leachman and Francis (2002) assess the long-run relationship between the two deficits. They suggest that higher budget deficits may have contributed to external deficits for the U.S. in the post-World War II period. Fidrmuc (2003) provides evidence for twin deficits in several countries, despite differences between the 1980s and the 1990s. Additionally, his study suggests that twin deficits emerged in the 1980s, though there is less evidence for twin deficits in the 1990s.

Third, some studies find evidence of a causality link between budget deficits and current account deficits, though the direction of causality is reversed (Anoruo and Ramchander, 1998; Khalid and Guan, 1999; Alkswani, 2000; Marinheiro, 2008; Kim and Kim, 2006). These studies argue that unidirectional causality exists and runs from current account deficits to budget deficits. The difference in the casual direction has arisen mostly from the character of countries being analyzed. Alkswani (2000) investigated the twin deficits hypothesis for a petroleum economy whose primary tax revenue is generated by oil. Anoruo and Ramchander (1998) studied whether the twin deficits hypothesis was valid for five developing Southeast Asian countries (India, Indonesia, Korea, Malaysia and the Philippines). They finds trade deficit to cause fiscal deficit and not vice versa for all the sample countries investigated, except Malaysia. Additionally, Anoruo and Ramchander (1998) argue that developing countries' macroeconomic dynamics influence budget deficits and current account deficits differently than in developed countries. Developed countries have large capital markets that provide a substantial portion of the funds needed for the financing of both public and private domestic needs. By contrast, many developing countries lack domestic capital markets (Kouassi et al., 2004; Anoruo and Ramchander, 1998).

Khalid and Guan (1999) find that current account deficits cause budget deficits in Indonesia and Pakistan. Marinheiro (2008) also finds evidence of this reverse but a weak long -run relationship between budget deficits and current account deficits. Marinheiro's study includes two explanations of this reverse causation. First, an increase in capital inflow leads to an appreciation of the currency, which causes a reduction in the current account. The reduction in the current account creates, *ceteris paribus*, a negative impact on domestic output, leading to decreased tax revenue and increased budget deficits. Second, the government could resort to fiscal incentives to diminish these effects. This scenario results in increased government expenditures and reduced tax revenues (Marinheiro, 2008).

When there is unidirectional causality from current account deficits to budget deficits, the government focuses on current account deficits. In other words, the government has the goal of eliminating current account deficits and uses budget deficits as a tool to achieve it (Marinheiro, 2008). In bidirectional causality, it does not suffice to decrease budget deficits to eliminate current account deficits (Marinheiro, 2008).

Fourth, some studies of this issue have concluded that the twin deficits hypothesis does not hold in developed countries, but that such a relationship exists in developing countries. According to Khalid and Guan (1999), this result stems from the problems afflicting developing countries, such as inefficient revenue collection systems and the lack of deep and sophisticated domestic capital markets to finance the budget deficit using domestic resources. Kouassi et al. (2004) reach similar results in their analysis, which suggests that the relationship between the two deficits exists in four out of five developing countries, while no developed country exhibits such a relationship. On the contrary, Chinn and Ito (2005) suggest that a balanced government budget has a positive and statistically significant impact on the current account of industrialized countries.

3. The Empirical Analysis

This section presents our empirical analysis. First, the data and methodology used in the paper are described, followed by a description of the results of testing for the twin deficit hypothesis. The long-run relationship between budget deficits and current account deficits can be representing by econometric model:

$$CAD_{it} = \alpha_i + \beta_i BD_{it} + \varepsilon_{it} \quad (10)$$

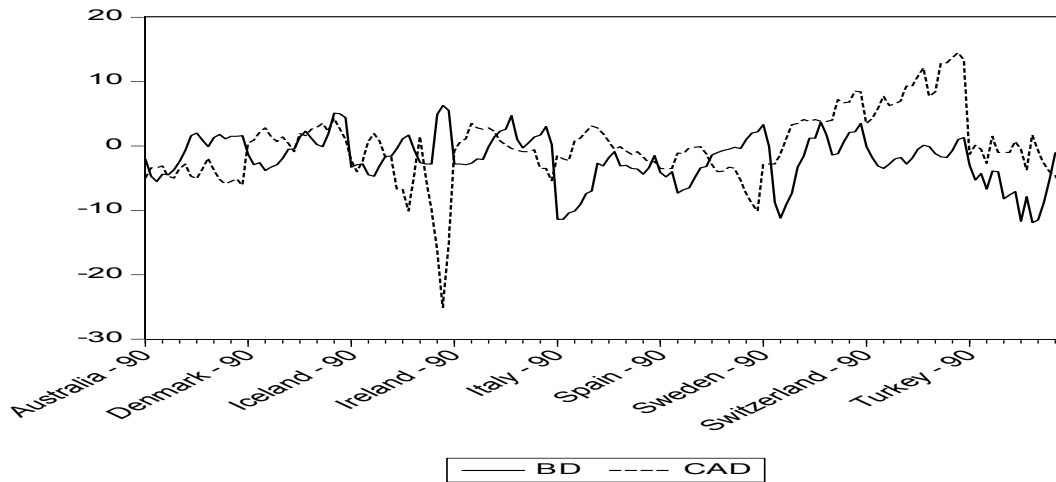
Where CAD_{it} and BD_{it} are, respectively, the current account and budget deficits, ε_{it} is error term. The existence of cointegration between CAD and BD is evidence that the two deficits do not move independently, as the REH perspective would have it.

3.1. Data

In testing for the relationship between budget deficits and current account deficits, we use the panel cointegration tests for nine OECD countries from 1990 to 2007. The main obstacle in this study is the lack of a long dataset for many countries. For this reason, we analyze a limited number of countries. We use annual data obtained from the OECD Factbook 2009 for the following countries: Australia, Denmark, Iceland, Ireland, Italy, Spain, Sweden, Switzerland, and Turkey, whose budget deficits have tended to diminish and current account deficits have increased, with the exceptions of Denmark, Sweden and Switzerland. The variables of interest in the study are budget deficits (BD) and current account deficits (CAD). The data on the budget deficits and current account deficits are measured as a percentage of GDP.

Figure 1 shows that there is a relationship between budget deficits and current account deficits for these nine countries from 1990 to 2007. Additionally, in the 1990s, budget deficits are a larger problem than current account deficits for all countries. In contrast with 1990s, since the 2000s, the current account imbalance has increased except in Denmark, Sweden and Switzerland. The significant increase in budget deficits in all countries in 1992 and 1993 stemmed from economic decline.

Figure 1: Budget Deficits and Current Account Deficits for OECD Countries (% GDP)



3.2. Methodology

Before applying the panel unit root and cointegration tests, we begin testing for cross-section dependence in the data, which is more reasonable. Then, the panel unit root and panel cointegration tests with structural breaks are employed.

3.2.1. Testing for Cross Section Dependence in Panel Data

This tests for cross section dependence can be carried out by using the Breusch and Pagan (1980) and Pesaran (2004) LM test statistics. The Breusch and Pagan LM test is based on the sum of squared coefficients of correlation among cross sectional residuals ($\hat{\varepsilon}_{it}$) obtained through OLS. The test statistics denoted by CD_LM_1 , CD_LM_2 , and CD_LM can be calculated as follows:

$$CD_LM_1 = T \sum_{i=1}^{N-1} \sum_{j=i+1}^N \hat{\rho}_{ij}^2 \tag{11}$$

$$CD_LM_2 = \sqrt{\frac{1}{N(N-1)}} \left(\sum_{i=1}^{N-1} \sum_{j=i+1}^N T \hat{\rho}_{ij} - 1 \right) \tag{12}$$

$$CD_LM = \sqrt{\frac{2T}{N(N-1)}} \left(\sum_{i=1}^{N-1} \sum_{j=i+1}^N \hat{\rho}_{ij} \right) \tag{13}$$

Where, the ρ_{ij} stands for the sample estimate of the cross sectional correlation among residuals. Under the null hypothesis of no cross sectional correlations, the CD_LM_1 statistic is distributed as chi-squared and CD_LM_2 and CD_LM test statistic is distributed as standard normal (Gülođlu and İvrendi, 2010).

3.2.2. Panel Unit Root Tests

Levin, Lin and Chu (2002) have developed a new panel unit root test with more power than standard ADF unit root tests. If the panel data frame y_{it} for $i=1,2,\dots,N$ and $t=1,2,\dots,T$ indicates the time series for individuals, and ADF (ρ) with an intercept can be shown as follows:

$$\Delta y_{i,t} = \alpha_i + \rho y_{i,t-1} + \sum_{j=1}^{p_i} \gamma_{ij} \Delta y_{i,t-j} + \xi_{i,t} \quad (14)$$

Where the test hypothesis is $\rho = 0$ versus $\rho < 0$ for $i=1,2,\dots,N$. Here, common unit root tests are applied for all cross-sectional individuals. However, Im, Pesaran and Shin (2003) (hereafter IPS) developed a panel unit root test which ρ can change for all cross-sectional individuals. We employ an IPS test with a null of $\rho_i = 0$ versus the alternative of $\rho_i < 0$ hypothesis for hysteresis for $i=1,2,\dots,N$.

$$\Delta y_{i,t} = \alpha_i + \rho_i y_{i,t-1} + \sum_{j=1}^{p_i} \gamma_{ij} \Delta y_{i,t-j} + \xi_{i,t} \quad (15)$$

Maddala and Wu (1999) propose the test statistic, which is based on combining the p-values of the test statistics (p_i) of N . Similar to IPS, this test allows for different first-order autoregressive coefficients and has the same null and alternative hypotheses in the estimation procedure. The test statistic (the Fisher test $P(\lambda)$) is as follows:

$$P(\lambda) = -2 \sum_{i=1}^N \ln(p_i) \quad (16)$$

where p_i is the p-value of the test statistic for unit i . The Fisher test statistic $P(\lambda)$ is distributed as a chi-squared distribution with $2N$ degrees of freedom (Esaka, 2003). In addition, MW statistics can be calculated similarly by combining the p-values of the test statistics of N independent Phillips-Perron (1988) regressions.

Hadri (2000) extends the tests of Kwiatkowski et al. (1992) to panel analysis. The null hypothesis is stationary in all individuals against the alternative of unit roots in all individuals. The test is constructed as a residual-based Lagrange multiplier test with the residuals taken from the regressions. The models may be expressed as

$$q_{it} = \beta_{irt}' d_{it} + \varepsilon_{it}, \quad i=1,\dots,N; \quad r=1,2 \quad (17)$$

where $\beta_{irt} = \beta_{i1t}$ when $r=1$ and $\beta_{irt} = (\beta_{i1t}, \beta_i)'$ when $r=2$. Hadri assumes that the variances of the ε_{it} are the same for every series. If the variances of the ε_{it} may not be the same for every series, the ε_{it} may be autocorrelated by considering the long-run variances of the ε_{it} and estimate them as

$$\hat{\sigma}_{\varepsilon_i}^2 = \frac{1}{T-1} \sum_{t=2}^T \hat{\varepsilon}_{it}^2 + 2 \sum_{j=1}^{\bar{k}} w_{kj} \left(\frac{1}{T-1} \sum_{t=j+2}^T \hat{\varepsilon}_{it} \hat{\varepsilon}_{i,t-j} \right) \quad (18)$$

The resultant statistic to test H_0 would be simply the average of the individual KPSS statistics for each series. Hadri shows that this statistic, appropriately standardized, will be asymptotically $N(0,1)$ under the null hypothesis (Erlat and Özdemir, 2003).

Pesaran (2007) proposes a panel unit root test for a balanced panel with N cross-sectional units and T time series observations. Pesaran (2007) considers the following heterogeneous, linear model:

$$\Delta y_{i,t} = \alpha_{0i} + \alpha_{1i}y_{i,t-1} + \gamma_i f_t + e_{it} \quad (19)$$

where f_t is an unobserved common factor, γ_i is the corresponding factor loading and e_{it} is an idiosyncratic error term independent across i and independent of the common factor. To account for the cross-sectional dependence included by the common factor, Pesaran (2007) suggests a cross-sectional augmentation of the test (Eq. (20)) with cross-sectional averages of the first differences and the lagged levels. The cross-sectional augmented (CA)DF equation is given by

$$\Delta y_{i,t} = a_i + b_i y_{i,t-1} + c_i \bar{y}_{t-1} + d_i \Delta \bar{y}_t + \varepsilon_{it} \quad (20)$$

where $\bar{y}_{t-1} = \sum_{i=1}^N y_{i,t-1}$, $\Delta \bar{y}_t = \sum_{i=1}^N \Delta y_{i,t}$ and ε_{it} is the regression error.

The individual specific test statistic for the hypothesis $H_{0i} : \rho_i = 1$ for a given i is now the t -statistic of b_i in Eq. (21), denoted by $CADF_i$. The panel unit root for the hypothesis $H_0 : \rho_i = 1$ for all i against the heterogeneous alternative $H_1 : \rho_i < 1$ for some i is given by the cross-sectional average of the $CADF_i$ tests as follows (Gengenbach, 2009):

$$\overline{CADF} = N^{-1} \sum_{i=1}^N CADF_i \quad (21)$$

The second test, presented by Pesaran (2007), refers to the inverse chi-squared, or Fisher, test (Eq. 14) and was developed by Maddala and Wu (1999). This test statistic uses a probability, defined in Eq. (21) by Maddala and Wu (1999). The third test, given Pesaran (2007), stands on the inverse normal test developed by Choi (2001). This test statistic can be calculated as follows:

$$Z(N, T) = \frac{1}{\sqrt{N}} \sum_{i=1}^N \Phi^{-1}(p_{iT}) \quad (22)$$

where $\Phi(\cdot)$ is the standard normal cumulative distribution function and this test statistic is also distributed as standard normal.

By adding the cross-section dimension, we increase the amount of information for each cross-section, thus solving the problems related to the lack of power of univariate unit root tests when the root is close to one, especially in small samples when the time dimension is restricted by the unavailability of long and reliable time series data. Moreover, using shorter samples with rich information helps us to avoid a second serious problem arising from the fact that standard unit root tests are biased towards the non-rejection of the null hypothesis in the presence of structural breaks. Obviously, as we reduce the sample length, the probability of discontinuities in the series generated either by shocks or by institutional changes diminishes (Gómez and Tamarit, 2011).

Although this first generation of tests is still being extensively used in the empirical literature, the main drawback (common to all them) is that they assume the absence of correlation across the cross-sections of the panel. That is, the individual members of the

panel (countries) are independent. This assumption is not realistic and, therefore, cannot be maintained in the majority of the cases, especially when the countries are neighbors or are involved in integration processes. A second generation of panel tests, by contrast, allows for different forms of dependence (see Pesaran, 2007), thereby eliminating the assumption of independence.

There are several alternative proposals formulated in the literature to overcome the cross-sectional dependency problem. First, Levin et al. (2002) suggest computing the test removing the cross-sectional mean. Although simple, this implies assuming, quite restrictively, that cross-sectional dependence is driven by one common factor with the same effect for all individuals in the panel data set. Second, Maddala and Wu (1999) propose obtaining the bootstrap distribution to accommodate general forms of cross-sectional dependence. Third, more recently, Pesaran (2007) suggests other proposals that are especially relevant when the dependence is pervasive, which is most commonly the case for integrated markets. Pesaran (2007) assumes that the process is driven by a group of common factors, so that it is possible to distinguish between the idiosyncratic component and the common component.

Levin et al. (2002), Im et al. (2003), Maddala-Wu (1999) and Hadri (2000) panel unit root tests ignore structural breaks. Therefore, a new panel data unit root test can be interpreted with structural breaks. This panel unit root test that includes a structural break is the PANKPSS test. The panel stationary test of Carrion-i-Silvestre et al. (2005) is a modification of Hadri's (2000) unit root test that allows for multiple structural breaks through the incorporation of dummy variables in the deterministic specification of the model. In this case, under the null hypothesis, the data-generating process (DGP) for the variable is assumed to be the following:

$$y_{i,t} = \alpha_i + \sum_{k=1}^{m_i} \theta_{i,k} DU_{i,k,t} + \beta_i t + \sum_{k=1}^{m_i} \gamma_{i,k} DT_{i,k,t}^* + \varepsilon_{i,t} \quad (23)$$

with dummy variable $DT_{ikt}^* = t - T_{b,k}^i$ for $t > T_{b,k}^i$ and 0 elsewhere, and another dummy variable $DU_{i,k,t} = 1$ for $t > T_{b,k}^i$ and 0 elsewhere, with denoting the k th date of the break for the i th individual, $k = \{1, 2, \dots, m_i\}$ and $m_i > 1$. The model in Eq. (24) includes individual effects such as individual structural break effects (shifts in the mean caused by the structural breaks), temporal effects if $\beta_i \neq 0$ and temporal structural break effects if $\gamma_{i,k} \neq 0$ (when there are shifts in the individual time trend).

According to Carrion-i-Silvestre et al. (2005), the specification given by Eq. (24) is general enough to allow for the following characteristics: (1) it permits individuals to have a different number of structural breaks; (2) the structural breaks may have different effects on each individual time series (the effects are measured by $\theta_{i,k}$ and $\gamma_{i,k}$); and (3) they may be located on different dates. The applied test of the null hypothesis of a stationary panel follows the test proposed by Hadri (2000), and the expression is given by:

$$LM(\lambda) = N^{-1} \sum_{i=1}^N \left(\hat{\omega}_i^{-2} T^{-2} \sum_{t=1}^T S_{i,t}^2 \right) \quad (24)$$

where $S_{i,t} = \sum_{j=1}^T \hat{\varepsilon}_{i,j}$ denotes the partial sum process that is obtained with the use of the estimated OLS residuals of (24), and where $\hat{\omega}_i$ is a consistent estimate of the long-run variance of $\varepsilon_{i,t}$. Additionally, the homogeneity of the long-run variance across individual time series can be imposed during the testing process. Finally, we use λ in Eq. (25) to denote the dependence of the test on the dates of the break. For each individual i , it is defined as the vector $\lambda_i = (\lambda_{i,1}, \lambda_{i,2}, \dots, \lambda_{i,m_i})' = (T_{b,1}^i / T, \dots, T_{b,m_i}^i / T)'$, which indicates the relative positions of the dates of the breaks during the entire time period T . Carrion-i-Silvestre et al. (2005) estimate the number of structural breaks and their positions by following the procedures put forth by Bai and Perron (1998) to compute the global minimization of the sum of the squared residuals.

3.2.3. Panel Cointegration Analysis

Larsson et al. (2001) suggest a maximum likelihood panel test of the cointegrating rank in heterogeneous panel models based on the mean of the individual rank trace statistics developed by Johansen (1995). For one individual, the trace statistics can be calculated as follows:

$$\lambda_{\text{trace}}(r) = -T \sum_{i=r+1}^p \ln(1 - \hat{\lambda}_i) \tag{25}$$

where $\hat{\lambda}_i$ is the i -th eigenvalue of the Π matrix. The LR_{bar} statistic is defined as the average of the N individual trace statistic as follows:

$$\overline{LR} = \frac{1}{N} \sum_{i=1}^N \lambda_{\text{trace},iT} \tag{26}$$

Larsson et al. (2001) propose to use a standardized LR_{BAR} statistic as a basis for the panel cointegration rank test defined by

$$LR_{\text{bar}} = \frac{\sqrt{N}(\overline{LR} - E(Z_k))}{\sqrt{\text{Var}(Z_k)}} \tag{27}$$

where $E(Z_k)$ is the mean and $\text{Var}(Z_k)$ is the variance of the asymptotic trace statistics. and values are tabulated by Larsson et al. (2001) in Table 1.

A Maddala-Wu (1999) test can also be used for the cointegration tests. The purpose of this test is to combine p -values from trace and max in the Johansen (1995) tests that are applied to each group in the panel data.

Pedroni (1999) proposes seven tests of the null hypothesis of no cointegration. These tests allow for homogeneity (Eq. (28)-(31)) and heterogeneity (Eq. (32)-(34)) in the long-run for the cointegrating vectors of each country in the panel. The process requires estimating the individual static cointegration regressions and then calculating from the estimated residuals the test statistics for the null hypothesis of no cointegration. Pedroni (1999) test statistics can be calculated as follows:

$$\text{Panel } \nu : Z_{\nu N,T} \equiv \left(\sum_{i=1}^N \sum_{t=1}^T L_{1li}^{-2} \hat{e}_{it-1}^2 \right)^{-1} \quad (28)$$

$$\text{Panel } \rho : Z_{\rho N,T-1} \equiv \left(\sum_{i=1}^N \sum_{t=1}^T L_{1li}^{-2} \hat{e}_{it-1}^2 \right)^{-1} \sum_{i=1}^N \sum_{t=1}^T L_{1li}^{-2} (\hat{e}_{it-1} \Delta \hat{e}_{it} - \hat{\lambda}_i) \quad (29)$$

$$\text{Panel PP: } Z_{tN,T} \equiv \left(\tilde{\sigma}_{NT}^2 \left(\sum_{i=1}^N \sum_{t=1}^T L_{1li}^{-2} \hat{e}_{it-1}^2 \right) \right)^{-1/2} \sum_{i=1}^N \sum_{t=1}^T L_{1li}^{-2} (\hat{e}_{it-1} \Delta \hat{e}_{it} - \hat{\lambda}_i) \quad (30)$$

$$\text{Panel ADF: } Z_{tN,T}^* \equiv \left(\tilde{S}_{NT}^{*2} \left(\sum_{i=1}^N \sum_{t=1}^T L_{1li}^{-2} \hat{e}_{it-1}^{*2} \right) \right)^{-1/2} \sum_{i=1}^N \sum_{t=1}^T L_{1li}^{-2} \hat{e}_{it-1}^* \Delta \hat{e}_{it}^* \quad (31)$$

$$\text{Group } \rho : \tilde{Z}_{\rho N,T-1} \equiv \sum_{i=1}^N \left(\sum_{t=1}^T \hat{e}_{it-1}^2 \right)^{-1} \sum_{t=1}^T (\hat{e}_{it-1} \Delta \hat{e}_{it} - \hat{\lambda}_i) \quad (32)$$

$$\text{Group PP: } \tilde{Z}_{tN,T} \equiv \sum_{i=1}^N \left(\tilde{\sigma}_{NT}^2 \sum_{t=1}^T \hat{e}_{it-1}^2 \right)^{-1/2} \sum_{t=1}^T (\hat{e}_{it-1} \Delta \hat{e}_{it} - \hat{\lambda}_i) \quad (33)$$

$$\text{Group ADF: } \tilde{Z}_{tN,T}^* \equiv \sum_{i=1}^N \left(\sum_{t=1}^T \tilde{S}_{NT}^{*2} \hat{e}_{it-1}^{*2} \right)^{-1/2} \sum_{t=1}^T \hat{e}_{it-1}^* \Delta \hat{e}_{it}^* \quad (34)$$

If there exist linear combinations of integrated variables that cancel out their common stochastic trends, then these series are said to be cointegrated. The economic translation is that these series share an equilibrium relationship. However, a commonly neglected phenomenon is that both the cointegrating vector and the deterministic components might change during the period analyzed, and if we do not take account of these structural breaks in the parameters of the model, then inference concerning the presence of cointegration can be affected by misspecification errors. In this case, the conventional tests suffer from a bias towards the spurious non-rejection of the null hypothesis of no cointegration. Therefore, in this paper we propose to use the tests developed in Banerjee and Carrion-i-Silvestre (2006), which generalize the approach in Pedroni (1999, 2004) to account for one structural break that may affect the long-term relationship in a number of different several ways (cointegrating vector and/or deterministic components).

We refer here to five of the cointegration models developed by Banerjee and Carrion-i-Silvestre (2006):

- Model 1: Constant term with a change in level but stable cointegrating vector

$$y_{it} = \mu_i + \theta_i DU_{it} + x'_{it} \delta_i + \varepsilon_{it} \quad (35)$$

- Model 2: Time trend with a change in level but stable cointegrating vector

$$y_{it} = \mu_i + \beta_i t + \theta_i DU_{it} + x'_{it} \delta_i + \varepsilon_{it} \quad (36)$$

- Model 3: Time trend with change in both level and trend but stable cointegrating vector

$$y_{it} = \mu_i + \beta_i t + \theta_i DU_{it} + \gamma_i DT_{it}^* + x'_{it} \delta_i + \varepsilon_{it} \quad (37)$$

- Model 4: Constant term with change in both level and cointegrating vector

$$y_{it} = \mu_i + \theta_i DU_{it} + x'_{it} \delta_{it} + \varepsilon_{it} \quad (38)$$

- Model 5: Time trend with change in both level and cointegrating vector (the slope of trend does not change)

$$y_{it} = \mu_i + \beta_i t + \theta_i DU_{it} + x'_{it} \delta_{it} + \varepsilon_{it} \quad (39)$$

where $DU_{it} = \begin{cases} 0 & t \leq T_{Bi} \\ 1 & t > T_{Bi} \end{cases}$, $DT_{it}^* = \begin{cases} 0 & t \leq T_{Bi} \\ (t - T_{Bi}) & t > T_{Bi} \end{cases}$, and $T_{Bi} = \lambda_i T$, $\lambda_i \in (0,1)$ denotes the

time of the break for the i -th country in the panel, $i=1,2,\dots,N$. We test the null hypothesis of no cointegration against the alternative hypothesis of cointegration, applying the ADF test statistics to residuals of the cointegration regression within the framework of the Pedroni (1999) test. Because the break point is unknown, Banerjee and Carrion-i-Silvestre (2006) propose to combine individual information in the panel data statistics. They calculate two test statistics, $Z_{\hat{t}_{NT}}$ and $Z_{\hat{\rho}_{NT}}$, as follows:

$$Z_{\hat{t}_{NT}}(\hat{\lambda}) = \sum_{i=1}^N t_{\hat{\rho}_i}(\hat{\lambda}_i) \quad (40)$$

and

$$Z_{\hat{\rho}_{NT}}(\hat{\lambda}) = \sum_{i=1}^N T \hat{\rho}_i(\hat{\lambda}_i) \quad (41)$$

where $Z_{\hat{t}_{NT}}(\hat{\lambda})$ denotes the pseudo t-ratio statistic, and $Z_{\hat{\rho}_{NT}}(\hat{\lambda})$ is the normalized bias. This strategy ensures a high degree of heterogeneity in the cointegrating vector. Additionally, these tests take into account differences in the short-run dynamics and the break point among individual countries in the panel. Using this panel cointegration approach allows us to increase the power of the statistical inference when estimating the parameters for individual countries in the panel.

3.3. Empirical Results

In light of all these explanations, we control for the following econometric issues usually neglected in earlier literature. First, we account for cross-sectional dependence among countries in the panel tests. Second, we allow for the existence of a break in the cointegration relationship, a major point to assess the effect of institutional changes in the relationship. Therefore, in this section, we present the results of cross-sectional dependence, panel unit root and panel cointegration tests.

3.3.1. Results of Cross-Section Dependence Tests

In this subsection, we test the null hypothesis of non-correlation against the alternative hypothesis of correlation using the approach suggested in Pesaran (2004). The test statistics with the corresponding probabilities are shown in Table 1. It is clear that the correlations among cross-sectional residuals do not show evidence of cross-sectional dependence according to the CD_{LM_1} and CD_{LM_2} tests. In contrast, the CD_{LM} test results show evidence of cross-sectional dependence for BD. However, this test is not valid when T is large and N is small, which is the case in our data. As a result, we may not allow for cross-sectional dependence when testing the unit root and cointegration of the series.

Table 1: Cross-sectional Dependence Tests Results

Test Statistics	BD	CAD
	Value (Prob.)	Value (Prob.)
CD_LM ₁	38.0762 (0.3751)	42.6664 (0.2063)
CD_LM ₂	0.2447 (0.4033)	0.7856 (0.2160)
CD_LM	2.4748 (0.0067)	1.2326 (0.1088)

3.3.2. Results of Panel Unit Root Tests

We should start the analysis by studying the order of integration of the variables. Several procedures to test for unit roots in panels are already available in the literature. We have applied the Fisher ADF test following Levin et al. (2002), Im et al. (2003), and Maddala and Wu (1999); the Fisher PP test of Maddala and Wu (1999) and Choi (2001); and the panel unit root tests of Hadri (2000) and Pesaran (2007), which ignore structural breaks.

Table 2: Panel Unit Root Test Results without Structural Breaks

Panel Unit Root Tests	BD	CAD
	Level	
Levin et al. (2002)	0.634	0.396
Im et al.(2003):W _{tbar} test	1.107	1.403
Fisher-ADF test (in Maddala and Wu, 1999; Choi, 2001)	8.694 & 1.249	11.131 & 1.549
Fisher-PP test (in Maddala and Wu, 1999; Choi 2001)	8.766 & 1.404	12.327 & 1.087
Hadri (2000):Homogeneous & Heterogeneous	4.716 ^{***} & 4.887 ^{***}	6.434 ^{***} & 5.548 ^{***}
Pesaran (2007): (in Im et al. 2003; Choi, 2001; Maddala and Wu, 1999)	-2.169 & -1.129 & 0.101	-1.947 & -0.632 & 0.264
	First Differences	
Levin et al. (2002)	-7.164 ^{***}	-7.341 ^{***}
Im et al.(2003):W _{tbar} test	-6.291 ^{***}	-7.119 ^{***}
Fisher-ADF test (in Maddala and Wu, 1999; Choi, 2001)	71.362 ^{***} & -6.065 ^{***}	81.029 ^{***} & -6.615 ^{***}
Fisher-PP test (in Maddala and Wu, 1999; Choi 2001)	70.661 ^{***} & -6.003 ^{***}	118.356 ^{***} & -8.066 ^{***}
Hadri (2000): Homogeneous & Heterogeneous	0.406 & 0.426	0.045 & 3.051 ^{**}
Pesaran (2007): (in Im et al. 2003; Choi, 2001; Maddala and Wu, 1999)	-3.79 ^{***} & -6.01 ^{***} & 0.00 ^{***}	-3.48 ^{***} & -5.09 ^{***} & 0.00 ^{***}

Note: Panel unit root test results according to model with intercepts. *** Significant at 1%, ** significant at 5%.

In Table 2, the null hypothesis of the Hadri test is rejected; in contrast, the null hypothesis in other panel unit root tests is not rejected. That is, the BD and CAD series has a unit root. When we take the first difference of series, the null hypothesis in the Hadri test is not rejected, but the null hypothesis in other tests is rejected. The results of the panel unit root test are affected by cross-sectional dependence, but our results are not changed for these two cases. Therefore, all of the panel unit root tests show the existence of unit roots in levels and no unit root in first differences of both BD and CAD. These results show that the BD and CAD series are $I(1)$, or integrated of order one.

The results in Table 3 illustrate the PANKPSS unit root test with multiple structural breaks, as proposed by Carrion-i-Silvestre et al. (2005). If we combine the individual information to compute the PANKPSS test in Table 3, the null hypothesis of stationarity can be rejected for both BD and CAD when the test is computed using homogeneous and heterogeneous long-run variance estimates. Therefore, it can be said that non-stationary elements of the series are not caused by structural breaks.

Table 3: Panel Unit Root Test Results with Structural Break

PANKPSS	BD	p-value	CAD	p-value
Homogeneous	8.4363***	0.0000	8.6154***	0.0000
Heterogeneous	32.5160***	0.0000	21.0936***	0.0000

Notes: *** Significant at 1%.

3.3.3. Results of Panel Cointegration Tests

The econometric methodology we use to analyze long-run relationships among the variables of our panel is based on cointegration techniques. These tests, originally applied and developed for time series, have been successfully adapted to the case of panel data. The main advantage of this methodology is that it overcomes the problem of the non-stationarity usually found in economic variables. The most common way to deal with this problem has been to take first differences. However, this filter removes from the variables an important part of the long-run information. Consequently, an alternative and more efficient way to test for long-run relationships in panels is to use tests for panel cointegration.

The BD and CAD series are integrated of order one, respectively. Therefore, we employ the panel cointegration tests of Pedroni (1999), Larsson et al. (2001) and Maddala-Wu (1999), which ignore structural breaks.

When we analyze the data in Table 4, all of the Pedroni test statistics exhibit no cointegration for BD and CAD. However, the tests of Larsson et al. and Maddala-Wu found a long-run relationship between BD and CAD. That is, these tests find evidence of the validity of the twin deficit hypothesis in the long run. Moreover, the existence of structural breaks in the cointegrating relationships biases the results in panel settings, as has been described in Banerjee and Carrion-i-Silvestre (2006), who propose an extension of the Gregory and Hansen (1996) approach. In addition, they use the common factors to account for dependence. When taking into consideration structural breaks, the Banerjee and Carrion-i-Silvestre (2006) test is applied to see whether these results have changed. Table 5 displays the results of the panel cointegration test with structural breaks.

Table 4: Panel Cointegration Tests Results without Structural Breaks

Cointegration tests	Test statistics	p-value
Pedroni (1999)		
Panel ν	0.4985	0.3523
Panel ρ	-0.2965	0.3818
Panel PP	-1.1591	0.2038
Panel ADF	-1.2714	0.1778
Group ρ	1.0228	0.2365
Group PP	-0.1151	0.3963
Group ADF	-1.4525	0.1389
Larsson et al. (2001)		
LR-bar $H_0: r=0, H_1: r>0$	26.239	0.0000
Maddala-Wu (1999)		
Trace $H_0: r=0, H_1: r>0$	38.45	0.0034
$H_0: r=1, H_1: r>1$	20.31	0.3159
Max $H_0: r=0, H_1: r>0$	34.11	0.0122
$H_0: r \leq 1, H_1: r>1$	20.31	0.3159

Notes: Panel unit root test results according to model with intercepts.

Table 5: The Banerjee and Carrion-i-Silvestre Panel Cointegration Tests Results with Structural Breaks

Cointegration Models	Test Statistics	p-value
Model 1:		
$Z_{\hat{I}_{NT}}(\hat{\lambda})$	-19.197	0.0000
$Z_{\hat{\rho}_{NT}}(\hat{\lambda})$	-5.379	0.0000
Model 2:		
$Z_{\hat{I}_{NT}}(\hat{\lambda})$	-11.104	0.0000
$Z_{\hat{\rho}_{NT}}(\hat{\lambda})$	-11.092	0.0000
Model 3:		
$Z_{\hat{I}_{NT}}(\hat{\lambda})$	-22.676	0.0000
$Z_{\hat{\rho}_{NT}}(\hat{\lambda})$	-10.329	0.0000
Model 4:		
$Z_{\hat{I}_{NT}}(\hat{\lambda})$	-21.732	0.0000
$Z_{\hat{\rho}_{NT}}(\hat{\lambda})$	-3.101	0.0009
Model 5:		
$Z_{\hat{I}_{NT}}(\hat{\lambda})$	-23.364	0.0000
$Z_{\hat{\rho}_{NT}}(\hat{\lambda})$	-18.311	0.0000

Notes: Panel unit root test results according to model with intercepts.

In Table 5, all models using the Banerjee and Carrion-i-Silvestre test statistics reject the null hypotheses that do not exhibit cointegration. Accordingly, the Banerjee and Carrion-i-Silvestre tests (Models 1-5) support the existing of a long-run relationship between BD and CAD in OECD countries. As a result, the Banerjee and Carrion-i-Silvestre test results support the twin deficit hypothesis for these countries.

4. Conclusion

This paper analyzes the validity of the twin deficit hypothesis in nine OECD countries from an empirical perspective. Our study is novel because it empirically examines the relationship between budget deficits and current account deficits using panel cointegration tests with structural breaks. First, we apply the test for cross-sectional dependence in the data because the results of the panel unit root test are affected by cross-sectional dependence. Then, the panel unit root tests are employed, both with and without structural breaks. We thereby take into account the impact of the economic crisis on budget deficits and current account deficits. Because of the unavailability of data for many countries, we include analysis for a limited number of countries. Through our analysis, we find a long-run relationship between budget deficits and current account deficits. The results of the panel cointegration tests, with and without structural breaks, show empirical evidence of the validity of the twin deficit hypothesis in the long run for selected countries between 1990 and 2007. The empirical results support the Keynesian proposition that there is a strong link between budget deficits and current account deficits. These findings support the previous results of Piersanti (2000), Rosenswieg and Tallman (1993), Normandin (1999) and Blanchard (1985).

These results suggest that government policy should be directed primarily toward the reduction of budget deficits in these OECD countries. In the Keynesian proposition, the impact of budget deficits on the economy is important. Reducing the budget deficit increases national savings, which reduces the current account deficit of a specific country. Solving the problem of budget deficits and current account deficits requires more of a focus on fiscal policy measures, such as tax increases or reductions in government expenditures, than monetary policy.

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