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A Panel Cointegration Analysis of the Relation between Private and Government Consumption

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21st December 2005

Abstract

This paper analyses the relation between private and government consumption in 23 OECD countries between 1970 and 2001. In particular it addresses the issue of whether government consumption is a substitute for or a complement to private consumption. The empirical analysis is made with panel cointegration analysis, using the newly developed CUSUM cointegration test by Westerlund (2005). The method is extended by using a bootstrap technique to control for cross-sectional dependence. The results show that government consumption is a complement to private consumption for most of the countries and a substitute for only a few of the countries.

JEL Classification: E21, E62, C33

Keywords: Private consumption, Government consumption, Panel cointegration

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1 Introduction

The aim of this study is to analyse the relationship between private and government consumption. Government spending exerts an important influence on aggregate demand, which in turn affects output, employment and consumption. The magnitude of this effect hinges partly on how private consumption reacts to changes in government spending. One channel of impact of government consumption on private consumption depends on whether these are substitutes or complements. For example, if the private sector is willing to substitute to a large extent their own expenditures for public expenditures, the effect of increases or decreases in government spending on aggregate demand and output is offset by a corresponding change in private consumption in the opposite direction. On the other hand, if government spending complements private consumption, aggregate demand unambiguously rises following an increase in government spending. In terms of so-called Edgeworth substitutes and complements, government consumption is said to be a substitute for private consumption if the marginal utility of private consumption decreases with government consumption and a complement if the marginal utility increases with government consumption. The composition of government consumption has a large influence on whether it is a complement to or substitute for private consumption. Some parts of government spending, such as defence and infrastructure, are probably complements to private spending, while other parts, such as spending on cultural activities and public transports, are substitutes. Many components of government spending also have both complementarity and substitutability effects on private consumption. Thus, the relationship has important implications for the design and understanding of macroeconomic and fiscal policy.

The question regarding the relation between private and government consumption has been addressed in several studies, both theoretically and empirically. The basis of these studies is often the concept of effective consumption, which is defined as the weighted sum of private and government consumption. The assumption is that it is effective consumption that enters the utility function and affects the marginal utility of the individuals, so that the private sector adjusts its level of consumption in connection with changes in government spending. Early studies in this area are Kormendi (1983) and Aschauer (1985), who find

that private and government spending are substitutes in the US economy. This hypothesis is supported by Ahmed (1986) and Katsaitis (1987) for the UK and Canada, respectively. Graham (1993) and Ni (1995) partly question this result by arguing that the relation between the variables is not robust to different empirical specifications. However, the general picture arising in these early studies is that private and government consumption can be regarded as substitutes.

The empirical analyses in these studies has been made with various statistical methods under the assumption that the variables included are stationary. However, economic variables such as private and government consumption often exhibit non-stationary characteristics. It is also reasonable to believe that the variables share some common stochastic trends and are thus cointegrated. Disregarding non-stationarity and cointegration may affect the results and implications drawn from the empirical analysis. To address these problems, some more recent studies have applied cointegration analysis to analyse the relation, e.g. Karras (1994), Henry and Olekalns (2001) and Amano and Wirjanto (1994). The results in these studies point in different directions; in some countries, private and government consumption are complements and in others they are substitutes. The results are often due to specific structural economic characteristics in the different countries. In addition to this, the results often depend on whether cointegration is found or not found in the empirical analysis. This mixed empirical result can be due to several circumstances. For example, cointegration tests often lack power, which makes it difficult to draw conclusions on statistical grounds. To have a broader base for the analysis, regarding both the underlying economic structures and the statistical techniques, panel model analysis has been introduced. Ho (2001) investigates the relation between private and government spending in a panel of OECD countries between 1981 and 1997, using panel cointegration analysis. He finds that government spending and private consumption can be regarded as substitutes, but in the statistical analysis he assumes that the exact relationship between the variables is the same for all 24 countries. This is, however, a questionable approach. Several structural factors in the economy, for example the size of the government sector and the composition of government spending, affect the relation between the variables and these factors certainly vary among countries.

In this study, the relation between private and government consumption in 23 OECD countries between 1970 and 2001 is analysed using panel cointegration analysis. We extend the analysis of Ho in a number of respects. First, we formally test if the relation between private and government consumption is the same for all countries. Our results show that the relation varies considerably among the countries. This is so regarding both the question of whether the variables are complements or substitutes, and the magnitude of the relationship. We find that private and government consumption are best described as complements in the majority of the OECD countries. Only in a few countries, do we find a significant relation of substitutability. However, if we incorrectly treat the countries as a homogenous unit and estimate an aggregate relation, this relation implies that the variables are substitutes. Hence, one has to be very cautious about analysing a group of countries in a uniform way. We also use a longer data set than Ho and a different test for panel cointegration. The cointegration test we apply is a CUSUM test for panel cointegration, newly developed by Westerlund (2005). The test is a residual-based test for cointegration that, like the vast majority of panel cointegration tests, assumes that there is no correlation between the members in the panel model. As this is a restrictive assumption in empirical applications, we conduct a bootstrap of the test, which allows us to model any cross-sectional dependence. The bootstrap results in substantially larger critical values for the test under the circumstances present in our study.

The paper is organized as follows. Previous research in this area is presented in more detail in section 2. In section 3, a simple theoretical model is derived, which serves as the basis for the empirical analysis. An overview of unit root tests and cointegration tests used in panel models is given in section 4. Here, we also discuss various estimation techniques for panel cointegration regressions, and present the CUSUM test later applied in the empirical analysis. A bootstrap analysis of the CUSUM test is made in section 5. The results of the cointegration analysis and estimations of the relation between private and government consumption are found in section 6. Conclusions are made in section 7.

2 Previous studies

The theoretical basis for the previous research of whether private and government consumption are substitutes or complements is often a model where effective consumption, instead of only private consumption, affects the utility of the private sector. The concept of effective consumption, first put forward by Bailey (1971), is that government expenditure on consumption goods and services adds to the welfare of the private households to the extent that the goods have a consumption value for them. Early macroeconomic studies often treated the government sector as separate from the private sector, as if households placed no value on the goods and services supplied by the government, and disregarded government expenditure entirely when making consumption decisions. The opposite scenario is that households, instead of disregarding the government sector, value government expenditure and fully regard government consumption as part of their total consumption. As Bailey argues, the relevant scenario may lie somewhere in between these two extremes. In that case, one unit of public consumption is valued by the private sector as much as θ units of private consumption, where $0 \leq \theta \leq 1$. Barro (1981) incorporates this idea in a theoretical model developed to study the output effects of government spending.

Kormendi (1983) and Aschauer (1985) study the US economy and find empirical support for the hypothesis that government consumption can be regarded as a substitute for private consumption. Kormendi estimates consumption functions with US data and finds a significant degree of substitutability between the variables. Aschauer asks to what extent government spending directly substitutes for private consumer expenditure in the private sector's utility function. He derives a theoretical model based on the permanent income hypothesis and estimates a relation derived from the Euler condition. He finds that public spending reduces private consumption expenditure on non-durable goods and services by 23 to 42 percent. The magnitude of the effect is compatible with Kormendi's results. Graham (1993) questions the results in Aschauer (1985) and claims that a suitable model for analysing the relationship between private and government consumption should include disposable income as an explanatory variable. The argument is that private consumption is found to track disposable income too closely for the permanent income hypothesis to be

consistent (Campbell and Mankiw, 1990). Extending the Aschauer model with disposable income, Graham obtains different results for the US example. The substitutability between private and government consumption is generally smaller in magnitude in Graham's model and for some empirical specifications the relation between the variables is the opposite, i.e. government consumption is a complement to private consumption. Graham also advocates analysing the question using a disaggregate measure of government spending.

Ni (1995) analyses how different specifications of the utility function affect the empirical estimations of the substitutability of government purchases for private consumption. The specifications Ni considers are the functional form of the utility function, the question of non-separability between non-durable and durable goods and services, the assumption of time separability of the utility function and the formal specification of effective consumption. Generally, the conclusion is that the results are sensitive to the choice of all those specifications. For example, when using an additive specification of effective consumption, government purchases are a substitute for private consumption, but with a Cobb-Douglas specification, government purchases are a complement to private consumption.

Ahmed (1986) and Katsaitis (1987) examine the relation between private and government consumption in the UK and Canada, respectively. Both authors find that government consumption is a substitute for private consumption, with a magnitude of approximately 40 percent for both countries. These results are in line with the results in Kormendi (1983) and Aschauer (1985) for the US.

In the above-mentioned studies, the questions of stationarity, non-stationarity and cointegration between the variables are not discussed to any greater extent. It is, however, reasonable to believe that private consumption, government consumption and disposable income are cointegrated. This hypothesis, in combination with the widespread use of cointegration techniques from the 1990s onwards, has led to several studies where cointegration analysis is used to empirically study the relation between private and government consumption.

Karras (1994) investigates the interaction between private and government consumption for a large number of countries. He tests for the presence of cointegration in the model,

but does not find any such evidence. He then analyses each country separately under the assumption of no cointegration. Karras' results contradict the earlier work by Kormendi (1983), Aschauer (1985) and Ni (1995). Private and government consumption are best described as complementary goods in the sense that government consumption tends to raise the marginal utility of private consumption. Substitutability is an exception and not the rule. Karras also finds that for countries with a smaller size of the government sector, the complementary relationship is stronger than in countries with a larger size of the government sector. In the latter countries, the relation between the variables moves towards substitutability.

Amano and Wirjanto (1994) focus on Canada and address the question of cointegration by performing cointegration tests in the framework of Engle and Granger (1987). The results are not conclusive as to whether cointegration is present or not, and the authors examine the question under two assumptions: 'cointegration' and 'no cointegration'. It turns out that the empirical results are highly dependent on this assumption. In a cointegration model, government consumption is a complement to private consumption, while in a model with no cointegration government consumption is a substitute for private consumption. In two papers, Amano and Wirjanto (1997, 1998) also look at the US economy using a cointegration model, and find private and government consumption to be substitutes. However, these papers are based on a theoretical model with the unappealing characteristic that each household individually can choose the amount of government consumption consumed. The same theoretical framework is used by Ho (2004) in a study of Japan.

Australia is examined by Henry and Olekalns (2001), where the Johansen framework for cointegration analysis (Johansen, 1988, 1991; Johansen and Juselius, 1990, 1992) is applied. The authors stress two aspects. They argue, first, that it is important to divide private consumption into consumption of durable and non-durable goods and services and, second, that one should take into account the presence of possible structural breaks in the model. Henry and Olekalns find that the relation between private and government consumption in Australia has changed over time. Before the 1980s, government consumption was a complement to private consumption while it has been a substitute since.

A mixed picture regarding the relation between private and government consumption arises from these single-country studies. Whether the variables are complements or substitutes often depends on whether or not cointegration is found in the models. In many studies, it is also hard to find evidence of cointegration. One reason for this can be that cointegration tests often lack power. This raises the possibility of examining the non-stationary data in a panel context instead, since this method often increases the power in the statistical methods used. A panel cointegration approach is used by Ho (2001), who analyses the relation between private and government spending using the panel cointegration test suggested by McCoskey and Kao (1998). Ho analyses 24 OECD countries between 1981 and 1997 and finds that government spending and private spending can be regarded as substitutes for each other.

3 Theoretical framework

The theoretical framework is based on a representative individual with preferences over private consumption and goods and services stemming from the government sector.¹ We assume that, in time period t , the individual maximizes her expected lifetime utility, defined as

$$V_t = E_t \left[\sum_{j=0}^{\infty} \psi^j u(c_{t+j}^*) \right]. \quad (1)$$

E_t is the conditional expectation operator based on information in period t , ψ is the individual's subjective discount factor and $u(\cdot)$ is a concave momentary utility function. c_t^* is the effective consumption, defined as

$$c_t^* = c_t + \theta g_t. \quad (2)$$

The effective consumption is the weighted sum of c_t , the real private consumption per capita, and g_t , the real government consumption per capita. θ measures the relation between private and government consumption, i.e. whether they are substitutes or complements to each other. The question of substitutability and complementarity is captured in

¹The theoretical framework in this section is used by Aschauer (1985), Amano and Wirjanto (1994), Karras (1994) and Ho (2001).

the cross second derivative of $u(c_t^*)$, denoted $u''(\cdot)_{cg}$. If $u''(\cdot)_{cg} < 0$, an increase in g_t reduces the marginal utility of c_t and the variables are Edgeworth substitutes, and if $u''(\cdot)_{cg} > 0$, an increase in g_t raises the marginal utility of c_t and the variables are Edgeworth complements (Ni, 1995). If the cross second derivative is equal to zero, the variables are Edgeworth independent. The specification of effective consumption in equation (2) implies that c_t and g_t are substitutes if $\theta > 0$ and complements if $\theta < 0$. To avoid a situation where a negative value of θ causes the utility function to be a decreasing function of g_t , which contradicts standard assumptions about utility functions, a function of g_t can be added to the utility function as

$$V_t = E_t \left[\sum_{j=0}^{\infty} \psi^j u(c_{t+j}^* + \phi(g_{t+j})) \right]. \quad (3)$$

For negative values of θ , a suitable choice of $\phi(\cdot)$ makes the marginal utility of g_t positive, as long as c_t is positive (Christiano and Eichenbaum, 1988). Assuming that $\partial\phi/\partial g > 0$ and that the individuals have no control over g_t , a maximization problem containing equation (3) can be solved ignoring the utility contribution from government consumption through the function $\phi(\cdot)$.

The individual maximizes her expected lifetime utility subject to a period-by-period budget constraint as

$$a_{t+1} = (a_t + y_t - c_t - \tau_t)(1 + r). \quad (4)$$

a_t is real financial assets, including government issued bonds, at the beginning of period t , y_t is real labour income and τ_t is tax payments. The real labour income, y_t , is exogenously given in the model. The variables are expressed in per capita terms. Forward substitution of equation (4) results in an expression of the budget restriction in terms of present discounted values as

$$\sum_{j=0}^{\infty} \left(\frac{1}{1+r} \right)^j E_t [c_{t+j}] = a_t + \sum_{j=0}^{\infty} \left(\frac{1}{1+r} \right)^j E_t [y_{t+j} - \tau_{t+j}] \quad (5)$$

where the present value of the consumption must be equal to the sum of the real wealth in period t and the present value of the disposable income. The government sector has a period-by-period budget constraint as

$$b_{t+1} = (b_t + g_t - \tau_t)(1 + r) \quad (6)$$

or

$$\sum_{j=0}^{\infty} \left(\frac{1}{1+r} \right)^j E_t [\tau_{t+j}] = b_t + \sum_{j=0}^{\infty} \left(\frac{1}{1+r} \right)^j E_t [g_{t+j}] \quad (7)$$

in terms of present discounted value.² b_t is government debt. According to equation (7), government tax receipts must be equal to the initial government debt added to the present value of government spending. It is assumed that the representative individual is forward looking and that she recognizes that a current increase of government debt implies future tax obligations. Hence, the budget restrictions of the individual and the government can be combined into an economy-wide budget restriction as

$$\sum_{j=0}^{\infty} \left(\frac{1}{1+r} \right)^j E_t [c_{t+j}^*] = (a_t - b_t) + \sum_{j=0}^{\infty} \left(\frac{1}{1+r} \right)^j E_t [y_{t+j}] + \sum_{j=0}^{\infty} \left(\frac{1}{1+r} \right)^j E_t (\theta - 1) [g_{t+j}] \quad (8)$$

stated in terms of effective consumption.

The problem facing the individual is now to maximize the expected lifetime utility in equation (3) subject to the budget restriction in equation (8). The Lagrangean equation for this problem is given by

$$\mathcal{L} = \sum_{j=0}^{\infty} \psi^j E_t [u(c_{t+j}^*)] - \lambda \left[\sum_{j=0}^{\infty} \left(\frac{1}{1+r} \right)^j E_t [c_{t+j}^* - y_{t+j} - (\theta - 1) g_{t+j}] - (a_t - b_t) \right] \quad (9)$$

where λ is the Lagrange multiplier. Along with the budget restriction in (8), the maximization problem yields a necessary first order conditions as

$$E_t [u'(c_{t+j}^*)] (\psi (1+r))^j = \lambda \quad \forall j = 0, 1, 2, \dots \quad (10)$$

Equation (10) can be used to derive the Euler equation for the individuals' consumption between two adjacent periods, t and $t+1$, as

$$u'(c_t^*) = \psi (1+r) E_t [u'(c_{t+1}^*)]. \quad (11)$$

The expression states that for the individual to choose an optimal time path for effective consumption, it cannot be the case that she could be better off by reducing effective

²The derivations of equations (5) and (7) have been made under the assumption that $\lim_{k \rightarrow \infty} \left(\frac{1}{1+r} \right)^k a_{t+k} = 0$ and $\lim_{k \rightarrow \infty} \left(\frac{1}{1+r} \right)^k b_{t+k} = 0$.

consumption in one period, t , and increasing it in period $t + 1$. The intertemporal rate of substitution has to be equal to the intertemporal rate of transformation. If we assume that the change in marginal utility from one period to the next is small, the time path of effective consumption can be approximated by

$$E_t [c_{t+1}^*] = [\psi(1+r)]^\sigma c_t^* \quad (12)$$

where $\sigma = -u'(c^*)/c^*u''(c^*)$ is the elasticity of the intertemporal substitution.³

Based on equation (12), an econometric relationship can be formulated as

$$c_{t+1}^* = \gamma c_t^* + \eta_{t+1} \quad (13)$$

where $\gamma = [\psi(1+r)]^\sigma$ and η_t is identically and independently distributed and satisfies $E_t[\eta_{t+1}] = 0$. An expression stated in terms of c_t and g_t instead of c_t^* can be obtained by substituting equation (2) into (13). This yields

$$c_{t+1} + \theta g_{t+1} = \gamma (c_t + \theta g_t) + \eta_{t+1}. \quad (14)$$

If $\gamma < 1$, c_t^* is a stationary, or an I(0), variable. If both c_t and g_t are I(1) variables, this suggests that if $\gamma < 1$, c_t and g_t are cointegrated, with θ being the cointegration parameter.

To empirically examine the relation between private and government consumption we can, based on equation (14), form an equation describing a long-run relationship between c_t and g_t and analyse the cointegration properties of this relation. The long-run relationship is

$$c_t = \alpha_0 - \theta g_t + v_t. \quad (15)$$

By applying appropriate cointegration tests and estimation methods, we can obtain an estimate of θ . In the empirical analysis, we will, however, take into account the arguments put forward by Graham (1993), that the relation between private and government consumption may be better examined in a model where disposable income has been included. We therefore extend the model with disposable income, y_t^d , and formulate the long-run cointegration relation as

$$c_t = \alpha_0 - \theta g_t + \vartheta y_t^d + v_t \quad (16)$$

³Expression (12) follows the result in Hall (1978), but for effective consumption instead of private consumption only.

This relation will be analysed in section 6, where a panel cointegration test and panel cointegration estimation techniques will be applied to the model, using data for 23 OECD countries.

4 Panel cointegration

In the last two decades, cointegration analysis has become an important econometric tool in macroeconomic applications. One reason for the widespread use of this method is that while many macroeconomic variables exhibit non-stationary behaviour, they also share common stochastic trends over time. Under these circumstances, statistical techniques for integrated data, such as unit root testing and cointegration analysis, are suitable for empirical analysis. Alongside the many empirical applications using these econometric tools, a wide range of econometric studies concerning unit roots and cointegration has emerged. Many econometric tests for unit roots, stationarity and cointegration exhibit unpleasant statistical characteristics such as size distortions and low power to detect false null hypotheses. In an attempt to reduce these problems, the interest in analysing non-stationary panel models, instead of single-equation time series models, has grown immensely in recent years. One motivation for this interest is that the cross-sectional dimension in the panel models can contribute information that, combined with the information in the time series dimension, helps us towards better inference regarding the underlying statistical properties of our model. The econometric research concerning non-stationary panel models can be divided into two categories. The first category focuses on testing for unit roots in the data, while the second category focuses on testing for cointegration. This section gives a short review of different approaches available when analysing non-stationary panels. The focus is on the issue of panel cointegration, but since unit root testing is also conducted in the empirical analysis, one part of the review is devoted to panel unit roots.⁴

⁴Overviews of the use of panel data models when analysing non-stationary data are found in, for example, Banerjee (1999) and Choi (2005).

4.1 Panel unit root tests

Two commonly used panel unit root tests are the tests proposed by Levin et al. (2002) and Im, Pesaran and Shin (2003) (the IPS test). The tests are based on an autoregressive model of a variable, y_t , as

$$\Delta y_{it} = \rho_i y_{it-1} + \sum_{j=1}^{k_i} \varphi_{ij} \Delta y_{it-j} + \mu_i + \delta_i t + u_{it} \quad (17)$$

where $i = 1, \dots, N$ is the cross-sectional dimension and $t = k_i + 2, \dots, T$ is the time series dimension. Both tests rely on the assumption that the individual panel members are independent of each other. The error term u_{it} is assumed to be independently normally distributed with mean zero and variance, σ_i^2 . μ_i and $\delta_i t$ capture deterministic components in the model, in terms of a constant and a time trend, respectively. The appropriate number of first difference terms, k_i , is assumed to be known. In practice, however, it can be estimated using, for example, information criteria. The null hypothesis of both tests is that $H_0: \rho_i = 0$ for all i , which implies that all individual series contain a unit root.

The alternative hypothesis of the test proposed by Levin et al. (2002) is $H_A: \rho_i = \rho < 0$, which implies that no individual series contains a unit root. The test is conducted by estimating equation (17) for each individual panel member to obtain an estimate of σ_i^2 . Each individual is then normalized with the standard deviation σ_i , the normalized data is pooled and equation (17) is estimated again, using the pooled data. Levin et al. (2002) then consider a t-test of the pooled estimator of ρ of the null hypothesis $H_0: \rho = 0$. In order to converge to a standard normal distribution, the pooled t-test should be mean and scale adjusted. One limitation of the test is the assumption that the ρ -parameter is the same for all individuals. The alternative hypothesis, of no unit root in any series in combination with the assumption of the same mean reversion process for all individuals, may also be too strong for some empirical applications.

The IPS test allows for more heterogeneity among the individuals. The alternative hypothesis of this test is formulated as $H_A: \rho_i < 0$ for $i = 1, \dots, n_1$ and $H_A: \rho_i = 0$ for $i = n_1 + 1, \dots, N$. The value of ρ_i is thus allowed to vary under the alternative hypothesis. The IPS test is performed by running individual ADF regressions, as in equation (17), and

calculating the individual t-statistics for testing $\rho_i = 0$. The IPS test statistic is then an average of the N individual statistics. This implies that the test concerns the significance of N independent unit root tests.

The assumption of independent panel members is a restrictive assumption when applying the tests in practice, and several tests have been proposed to overcome this problem. Maddala and Wu (1999) suggest a new test for unit roots that allows for unbalanced panels and more heterogeneity among the individuals. By using a bootstrap technique, Maddala and Wu also allow the cross-sectional dependency between the panel members to take a more general form. Other tests that allow for cross-sectional dependency have been developed by Bai and Ng (2004) and Moon and Perron (2004) under the assumption that the correlation between the individuals can be modelled by common factors. Pesaran (2005) approaches the cross-sectional dependency differently. Individual ADF regressions are augmented with the cross-section averages of the lagged levels and first differences of the individual series, which remove the cross-sectional correlation. Based on these augmented ADF statistics, Pesaran develops a modified version of the IPS test and provides critical values for it.

Hadri (2000) takes another approach, by proposing a test with the null hypothesis of stationarity for all panel members. The alternative hypothesis is that each series contains a unit root. The test by Hadri can be seen as a panel average of the test for level stationarity proposed by Kwiatkowski et al. (1992).

The IPS test and the test proposed by Levin et al. (2002) have often been used in practice in empirical applications. However, economic panels are seldom cross-sectionally independent and the assumption of no cross-sectional correlation is very restrictive. It is therefore wise to apply a unit root test that allows for this kind of correlation.

4.2 Panel cointegration regressions

The majority of statistical tests for cointegration in panel data models are based on estimated residuals from a panel cointegration regression.⁵ To examine the cointegration properties, a panel test for unit roots or stationarity is then applied to the residuals. A panel cointegration model in this context can be formulated as

$$y_{it} = \alpha_i + \delta_i t + x'_{it} \beta_i + u_{it} \quad (18)$$

$$x_{it} = x_{i,t-1} + \nu_{it} \quad (19)$$

for $i = 1, \dots, N$ and $t = 1, \dots, T$. x_{it} is an $(M \times 1)$ -vector containing the regressors as $x_{it} = (x_{1it}, x_{2it}, \dots, x_{Mit})'$ and β_i is a corresponding parameter vector as $\beta_i = (\beta_{1it}, \beta_{2it}, \dots, \beta_{Mit})'$. α_i and $\delta_i t$ are deterministic components in terms of a constant and a time trend, respectively. It is assumed that y_{it} and x_{it} are $I(1)$ and that there is no cross-sectional correlation among the panel members. The error terms in the cointegration regression, u_{it} , and the first difference of x_{it} , ν_{it} , can be combined in a matrix as $w_{it} = (u_{it}, \nu'_{it})$. The long-run covariance matrix of w_{it} is defined as

$$\Omega_i = \sum_{j=-\infty}^{\infty} \Gamma_i(j) = \begin{pmatrix} \omega_{i11}^2 & \omega'_{i21} \\ \omega_{i21} & \Omega_{i22} \end{pmatrix} \quad (20)$$

where $\Gamma_i(j) \equiv E(\omega_{it} \omega'_{it+j})$ is the j^{th} order autocovariance of w_{it} . The long-run variance of u_{it} conditional on ν_{it} can be defined as

$$\omega_{i1.2}^2 \equiv \omega_{i11}^2 - \omega'_{i21} \Omega_{i22}^{-1} \omega_{i21}. \quad (21)$$

If y_{it} and x_{it} are cointegrated, u_{it} and ν_{it} are generally correlated in the long-run. If u_{it} and ν_{it} are correlated, ω_{i21} will not be zero, which causes an OLS estimator applied to equation (18) to have a second-order bias. In the presence of endogeneity between y_{it} and x_{it} and serial correlation in u_{it} , the OLS estimator becomes both biased and inefficient. As a consequence, statistical inference of cointegration tests based on OLS residuals becomes

⁵Tests for panel cointegration based on VAR models are proposed by Larsson et al. (2001) and Groen and Kleibergen (2003). Since this study uses a residual-based cointegration test, the VAR alternatives are not presented in this section.

problematic, since the distribution of the cointegration tests may depend on nuisance parameters. As a remedy for these problems, different modifications of the OLS estimator have been proposed, which are unbiased and asymptotically efficient even under the presence of endogeneity and serial correlation. Such estimators are the dynamic OLS (DOLS) proposed by Saikkonen (1991) and Stock and Watson (1993) and the fully-modified OLS (FMOLS) proposed by Phillips and Hansen (1990).

The DOLS approach is based on transforming the regression in equation (18), in order to eliminate the effects of the correlation between u_{it} and ν_{it} . Under the assumption of cointegration between y_{it} and x_{it} , both u_{it} and ν_{it} are stationary and a linear projection of u_{it} on all lags and leads of ν_{it} can be made in the form of

$$u_{it} = \sum_{j=-q}^q \xi'_{ij} \nu_{it} + u_{it}^*. \quad (22)$$

u_{it}^* is an orthogonal error term. With equation (22) in hand, equation (18) can be rewritten as

$$y_{it} = \alpha_i + \delta_i t + x'_{it} \beta_i + \sum_{j=-q}^q \xi'_{ij} \Delta x_{it} + u_{it}^* \quad (23)$$

since $\nu_{it} = \Delta x_{it}$. Assuming that the correlation between ν_{it+j} and u_{it} is equal to zero for $|j| > q$, equation (23) can be estimated with q lags and leads of Δx_{it} . In practice, the appropriate number of q can be chosen with information criteria. The DOLS estimator will be asymptotically unbiased and u_{it}^* can be used in residual-based cointegration tests without problems concerning nuisance parameters.

The FMOLS estimator is constructed by making corrections for the endogeneity and the serial correlation directly to the OLS estimator. To correct for the endogeneity problem, a consistent estimator of ω'_{i21} and Ω_{i22} , $\hat{\omega}'_{i21}$ and $\hat{\Omega}_{i22}$ say, can be used to modify equation (18) as

$$y_{it}^* = y_{it} - \hat{\omega}'_{i21} \hat{\Omega}_{i22} \nu_{it} = \alpha_i + \delta_i t + x'_{it} \beta_i + u_{it} - \hat{\omega}'_{i21} \hat{\Omega}_{i22} \nu_{it} \quad (24)$$

A correction for the serial correlation problem is achieved with the help of a consistent estimator of $\Lambda_i \equiv \sum_{j=0}^{\infty} \Gamma_i(j)$, which is the one-sided long-run covariance matrix of w_{it} . Details about the FMOLS estimator, as well as the DOLS estimator, can be found in Kao and Chiang (2000). Kao and Chiang analyse and compare the OLS, DOLS and

FMOLS estimators and show that all estimators are asymptotically normally distributed. Since DOLS and FMOLS are asymptotically equivalent, which of them to use in practice depends on their small sample properties. Based on the results from a Monte Carlo study, Kao and Chiang argue that the DOLS estimator is preferable to the FMOLS estimator.

4.3 Panel cointegration tests

The residual-based tests for panel cointegration can be divided into two categories: tests with the null hypothesis of no cointegration and tests with the null hypothesis of cointegration. Kao (1999) proposes tests with the null hypothesis of no cointegration. The alternative hypothesis in these tests is the presence of one common cointegration relationship between the variables that is valid for all panel members. Kao applies various versions of the Dickey–Fuller test and an augmented Dickey–Fuller test to the residuals from a panel estimation of equation (18) where the cointegration parameters in β are assumed to be equal across all individuals. Pedroni (1999) allows for heterogeneous cointegration relations and relaxes the assumption of equal β -parameters. He proposes seven tests of the null hypothesis of no cointegration, by constructing panel versions of the Phillips–Perron test and the augmented Dickey–Fuller test and applying them to residuals from a panel estimation. Pedroni divides the tests into two categories. In the first category, the so-called panel tests, the alternative hypothesis is that all individuals are cointegrated and, in the second category, the group mean tests, a subset of the panel members is cointegrated while the other subset is not.

While most cointegration tests have a null hypothesis of no cointegration, McCoskey and Kao (1998) consider a test with the opposite null hypothesis. A priori, economic theories often predict cointegration relations among economic variables and, therefore, it has been argued that it would be more natural to consider the null hypothesis of cointegration instead of the null of no cointegration in a statistical test. The test by McCoskey and Kao is also based on residuals from the estimation of the cointegration relation in equation (18). The cointegration parameters in β are allowed to vary between individuals but the test has an alternative hypothesis implying that there is no cointegration in any panel member.

Endogeneity and serial correlation under the null hypothesis require that the residuals are estimated with an unbiased and asymptotically efficient estimator such as the DOLS or the FMOLS estimator. The test is an LM test, constructed in the same framework as the time series stationarity test by Kwiatkowski et al. (1992) and the cointegration test by Shin (1994), as well as the panel stationarity test by Hadri (2000).

Westerlund (2005) proposes a residual-based CUSUM test of the null hypothesis of cointegration. Unlike the LM test by McCoskey and Kao, the CUSUM test allows for a heterogeneous alternative hypothesis, where some panel members are cointegrated while others are not. Further, the CUSUM test has been shown to have better properties regarding empirical size and size distortions than the LM test, if the residuals exhibit serial correlation (Westerlund, 2005). Since many empirical macroeconomic models exhibit serial correlation, this characteristic is of great importance. The CUSUM test is applied in the empirical analysis in this study and is presented in detail in the following section.

4.4 The CUSUM test

In the empirical analysis, we use the CUSUM test for panel cointegration proposed by Westerlund (2005).⁶ The null hypothesis of the test is that all panel members are cointegrated, while the alternative hypothesis allows for a mixture of cointegrated and non-cointegrated panel members. The idea behind the test is that if a set of variables is cointegrated, the residuals from a cointegration regression are stationary and, in that case, the residual series should be stable and not display large fluctuations. On the other hand, if the variables are not cointegrated, the residual series is a unit root process and is expected to fluctuate to a greater extent. The null hypothesis of the CUSUM test tests the degree of fluctuation in the residual series. If the residuals fluctuate too much, the null hypothesis of cointegration is rejected.

The CUSUM test is based on the panel cointegration model in equations (18) and (19). The test statistic is a cross-sectional average of the test statistic of the time series

⁶The author would like to thank Joakim Westerlund for providing programme codes for the application of the CUSUM test.

test proposed by Xiao and Phillips (2002). The test allows for heterogenous cointegration vectors and the cointegration regression is estimated separately for each individual. The DOLS or the FMOLS estimator has to be applied to estimate an asymptotically efficient residual series, so that the CUSUM test statistic is asymptotically free from nuisance parameters. The test statistic is calculated as

$$CS_{NT} = \frac{1}{N} \sum_{i=1}^N \left(\max_{t=1, \dots, T} \frac{1}{\hat{\omega}_{i,1.2} T^{1/2}} |S_{it}^*| \right) \quad (25)$$

where $S_{it}^* = \sum_{j=1}^t \hat{u}_{ij}^*$ is the cumulative sum of residuals and $\hat{\omega}_{i,1.2}$ is a consistent estimator of the long-run variance between u_{it} and Δx_{it} in equation (21). The statistic measures the magnitude of the residual variation from the regression of y_{it} on x_{it} against the magnitude of the estimated long-run conditional variance of u_{it} given ν_{it} . If y_{it} and x_{it} are cointegrated, the statistic should stabilize, but if they are not, the residual variation will increase and cause the statistic to diverge (Westerlund, 2005). The long-run covariance matrix, Ω_i , is estimated using a semiparametric kernel function as

$$\hat{\Omega}_i = \sum_{j=-M}^M \omega(j/M) \hat{\Gamma}_i(j) \quad (26)$$

where $\hat{\Gamma}_i(j) = T^{-1} \sum_{t=j+1}^T \hat{w}_{it} \hat{w}'_{it+j}$ and $\hat{w}_{it} = (\hat{u}_{it}^*, \hat{\nu}'_{it})'$. $\omega(j/M)$ is a kernel function which depends on the bandwidth parameter M . In the estimations, we apply the Bartlett kernel function $\omega(j/M) = 1 - j/(1 + M)$.⁷ The choice of M is important, since an inappropriate M may lead to inconsistency of the CUSUM test. Here, we choose $M = T^{1/3}$.⁸

By standardizing the CUSUM test by the first two moments, μ and σ^2 , of the asymptotic distribution of CS_{NT} as

$$Z_{NT} = \frac{N^{1/2}(CS_{NT} - \mu)}{\sigma}, \quad (27)$$

Westerlund (2005) shows that Z_{NT} converges to a standard normal distribution and also provides appropriate first and second moments. The moments depend on the deterministic components that are included in the model and on the number of regressors. The test is

⁷The Bartlett kernel has been found to perform best for the CUSUM test (Westerlund, 2005).

⁸Andrews (1991) shows that the optimal bandwidth for the Bartlett kernel applied to stationary time series is $M = O_p(T^{1/3})$.

one-sided, and the null hypothesis is rejected if the test statistic exceeds the critical value corresponding to the chosen significance level.

5 A bootstrap of the CUSUM test

The distribution of a statistical test is often derived on the basis of a variety of assumptions regarding the underlying data generating process. Some of these assumptions may not hold in practice when the tests are applied to true data series in econometric applications. One of several problematic consequences this can have for a test is size distortions, i.e. that the empirical size is larger than the nominal one. To overcome such a problem, one way is to bootstrap the test using an appropriate bootstrap method.⁹

In this section, we conduct a bootstrap of the CUSUM test for panel cointegration. The aim is to generate bootstrap critical values for the test that are more appropriate to use under the circumstances present in this study. There are some reasons to suggest that the use of a bootstrap should improve the properties of the test. The CUSUM test relies on the assumption of no cross-section correlation among the individuals.¹⁰ Violation of this assumption is one factor that may give rise to size distortions. A bootstrap allows us to model any cross-sectional dependence in the data. Moreover, the standardized test statistic of the CUSUM test is shown to converge to a standard normal distribution, a convergence that depends on the use of appropriate first and second moments of the test. Westerlund (2005) provides both asymptotic moments as well as finite sample moments, which are approximated with means and variances from a Monte Carlo simulation. One question is how fast the convergence to the normal distribution is in finite samples, a question that is related to the appropriateness of the approximated moments. If the approximated moments are inappropriate in a specific empirical application, this can also lead to a bias

⁹A brief overview of bootstrap methods is found in Bergström (1999). The use of bootstrapping in cointegration analysis is discussed in Li and Maddala (1997).

¹⁰To our knowledge, there are no residual-based tests, for panel cointegration with the null hypothesis of cointegration, available today, that allow for cross-sectional correlation. The VAR based cointegration test proposed by Groen and Kleibergen (2003) does, however, allow for such correlation.

in the test. Finally, when applying the CUSUM test, we have to estimate the panel cointegration model with the DOLS or the FMOLS estimation method to get unbiased and asymptotically efficient residuals. With the DOLS estimator, we add q lags and leads of Δx_{it} to the regression equation to account for the correlation between u_{it} and ν_{it} . If the number of lags and leads is not correctly chosen, this will be a source of bias as well.

The bootstrap is based on the panel cointegration model in equations (18) and (19) using data on private consumption, government consumption and income for the 23 OECD countries in the study.¹¹ The variable y_{it} is represented by private consumption and the two-dimensional variable vector x_{it} consists of government consumption and gross domestic product.¹² Deterministic components are included in the form of a constant. Under the null hypothesis of the CUSUM test, i.e. cointegration among the panel members, the model provides us the information that x_{it} and y_{it} are I(1) variables and that equation (18) is a cointegration relation. Under these circumstances, Li and Maddala (1997) propose the following bootstrap strategy. First, estimate equation (18) with an appropriate estimation method to get estimates of β_i and u_{it} and, second, define $\nu_{it} = \Delta x_{it}$ and bootstrap the pairs (u_{it}, ν_{it}) . The method preserves the correlation between u_{it} and ν_{it} , i.e. the endogeneity between y_{it} and x_{it} . In order to also take into account the structure of serial correlation in the errors, a moving block bootstrap is suggested. This general bootstrap structure is applied in the following bootstrap of the CUSUM test. More specifically, the bootstrap is conducted as follows:

1. With the original data and sample period in hand, equation (18) is estimated for each country separately. The DOLS and the FMOLS estimators are applied to account for endogeneity and serial correlation in the model.
2. The estimated cointegration parameters $\hat{\alpha}_i$ and $\hat{\beta}_i$ are saved in a $(3 \times N)$ matrix, $\hat{\beta}$. The estimated individual residual series $\hat{u}_{it} = \hat{\alpha}_i - x'_{it}\hat{\beta}_i$ is then extracted and

¹¹The data is described in detail in section 6.1. The sample period is 1970 to 2001. The data is on a quarterly basis and collected from OECD Economic Outlook No. 75.

¹²Unfortunately, data on disposable income is not available for all countries and we use GDP as a measure of income instead.

collected into a $(T \times N)$ matrix, \bar{u}_t . $\nu_{it} = \Delta x_{it}$ is calculated for each country and gathered into a matrix as $\bar{v}_t = (\Delta x_{1it}, \Delta x_{2it})$. The matrix of error terms, $\bar{w}_t = (\bar{u}_t, \bar{v}_t)$, contains the estimated residuals as well as the first differences of x_{it} .

3. Blocks of rows in \bar{w}_t are drawn with replacement and resampled to build up a new matrix of error terms for the bootstrap estimations. To account for the structure of serial correlation in the errors, a moving block bootstrap is used. The T observations are divided into blocks of length ℓ , where the blocks of observations may overlap each other. We conduct a set of bootstraps where the length of the moving blocks are varied from $\ell = 1, \dots, 8$. By doing this, we can trace the effect of allowing the serial correlation to span longer time horizons. Note that setting $\ell = 1$ leads to a non-block bootstrap in that the rows in \bar{w}_t are resampled one by one. By bootstrapping the pairs (\bar{u}_t, \bar{v}_t) in \bar{w}_t , the endogeneity between y_{it} and x_{it} is preserved.
4. New data series are constructed using the randomly drawn error terms from \bar{w}_t and the parameter estimates in $\bar{\beta}$. The bootstrap counterpart to x_{it} , \bar{x}_{it} , is created recursively as $\bar{x}_{it} = \bar{x}_{it-1} + \hat{u}_{it}$ with initial values set to zero and \bar{y}_{it} is created based on \bar{x}_{it} as $y_{it} = \hat{\alpha}_i + \bar{x}'_{it}\hat{\beta}$.
5. The model in (18) is estimated with the resampled data using the DOLS and the FMOLS estimator. The panel CUSUM test is applied to the estimated residuals and the test statistic is saved. The resample procedure is repeated 500 times, yielding 500 CUSUM statistics saved.
6. Finally, the 90, 95 and 99 percentiles of the CUSUM test statistics are extracted. These will be compared to the analytical critical values for the 10, 5 and 1 percent significance levels, respectively.

The results of the bootstrap are shown in table 1. In panel A, the analytical critical values based on the standard normal distribution are shown while the critical values from the bootstrap are presented in panel B. We distinguish between the DOLS and the FMOLS estimator and present critical values for different lengths of the blocks of residuals, ℓ .¹³

¹³To save space, only the results for $\ell = 1, 4, 5$ are presented in the table.

Table 1: Bootstrap of the CUSUM test

Panel A: Analytical critical values				
		10%	5%	1%
		1.28	1.64	2.33
Panel B: Bootstrapped critical values				
		10%	5%	1%
DOLS	$\ell = 1$	3.45	4.04	5.46
	$\ell = 4$	4.34	5.13	6.45
	$\ell = 5$	4.90	5.51	7.01
FMOLS	$\ell = 1$	2.33	2.79	3.67
	$\ell = 4$	3.30	3.83	4.95
	$\ell = 5$	3.48	4.08	5.37

Note: The analytical critical values are from the standard normal distribution. ℓ is the length of the moving blocks.

$\ell = 1$ implies that a non-block bootstrap is undertaken and that the error terms are drawn individually. In our empirical analysis, we use quarterly data and, therefore, a time span of the serial correlation of 4 to 5 periods may be reasonable.

The general conclusion is that the bootstrapped critical values are larger than the analytical ones and that the differences are substantial. The 5 percent critical value from the standard normal distribution is 1.64, which can be compared to 4.04 and 2.79 for DOLS and FMOLS respectively, when setting $\ell = 1$. With $\ell > 1$, i.e. when a moving-block bootstrap is made, the critical values are even larger. That the bootstrap would generate larger critical values than the ones from the standard normal distribution are in line with our expectations and our previous argumentation. For example, cross-sectional correlation in the panel model would cause size distortions and over-rejections of the test, which requires larger critical values. Generally, using the DOLS estimator results in larger critical values than using the FMOLS estimator. The overall conclusion from the bootstrap is that the differences in critical values, compared to the analytical ones, are substantial, which indicates that it is important to take into account special circumstances in an empirical application of the CUSUM test.

6 Empirical Analysis

6.1 Data and empirical model

The empirical analysis aims to examine the relation between private and government consumption by estimating a cointegration model for a panel of 23 OECD countries.¹⁴ The theoretical framework in section 3 serves as the basis for the analysis. The relation that is to be estimated is equation (16), where private consumption, government consumption and income form a cointegration relation. The long-run relation between private and government consumption in equation (15) is thus extended with a measure of income, in line with the arguments in Graham (1993). Unfortunately, due to lack of data, we have to use GDP as the measure of income instead of disposable income. Data on private consumption, government consumption and GDP has been collected from OECD Economic outlook No. 75. We employ data measured on a quarterly basis and all variables are in real per capita terms and expressed in terms of the US dollar PPP exchange rate of 2000. The sample period is 1970Q1 to 2001Q4.

Recapitulating the cointegration relation of interest, the specific empirical relation we want to estimate is

$$pc_{it} = \alpha_i + \beta_{1i} gc_{it} + \beta_{2i} gdp_{it} + \epsilon_{it} \quad (28)$$

where pc_{it} is private consumption, gc_{it} is government consumption and gdp_{it} is income measured as GDP. The empirical model includes deterministic components in the form of a country-specific constant term. The parameter β_1 is related to θ , the theoretical parameter describing the relation between private and government consumption, as $\beta_1 = -\theta$. Hence, if $\beta_1 > 0$, government consumption is a complement to private consumption, while it is a substitute if $\beta_1 < 0$.

¹⁴The countries are Australia, Austria, Belgium, Canada, Switzerland, Germany, Denmark, Spain, Finland, France, United Kingdom, Greece, Ireland, Iceland, Italy, Japan, Luxembourg, Netherlands, Norway, New Zealand, Portugal, Sweden and United States.

Table 2: Test of panel unit roots

		Levels		First diff.
Im, Pesaran and Shin (2003)	pc_t	7.51	Δpc_t	-13.36
	gc_t	4.51	Δgc_t	-15.22
	gdp_t	9.49	Δgdp_t	-15.15
Pesaran (2005)	pc_t	-0.38	Δpc_t	-2.85
	gc_t	-0.18	Δgc_t	-2.91
	gdp_t	-0.05	Δgdp_t	-3.04

Note: The 5 percent critical value for the IPS test is -1.64 and for the Pesaran test -2.20.

6.2 Unit root test

To examine the cointegration properties of the model, we apply the panel CUSUM test, described in section 4.4. For the test to be appropriate, we have to assume that the three variables are all I(1). To verify this assumption, we test for the presence of unit roots in the variables with a panel unit root test. We apply two panel unit root tests, first, the IPS test proposed by Im et al. (2003) and, second, the test proposed by Pesaran (2005).¹⁵ (See section 4.1 for a brief description.) While the IPS test assumes no cross-sectional dependence among the panel members, the Pesaran test allows for this kind of correlation. The null hypothesis of the tests is that all individual series contain a unit root, and, under the alternative hypothesis, the tests allow for both unit roots and stationarity among the individuals. The results of the two tests are presented in table 2.¹⁶

The general conclusion from the unit root tests is that all three variables can be regarded as being I(1) variables. The null hypothesis of a panel unit root is not rejected for the variables in levels, while it is rejected, in favour of stationarity, for the first differences. This conclusion can be drawn from both tests. The large test statistics, in absolute terms, for the IPS test are a result of the fact that the IPS test assumes no cross-section dependence in the panel, an assumption that is not valid in this application.

¹⁵The results of the individual ADF regressions underlying both the IPS test and the Pesaran test are presented in table 6 in the appendix. The results of the tests imply that the vast majority of the individual series appear to be I(1).

¹⁶The lag lengths in the underlying ADF regressions have been selected as the integer of $T^{1/3}$, where T is the number of observations. The choice of lag length does not influence the outcome of the test.

6.3 Cointegration results

In this section, we analyse the cointegration properties of the empirical model in equation (28). First, we apply the CUSUM test to the panel of the 23 OECD countries to test for panel cointegration. After finding support for the hypothesis of cointegration, we analyse the outcome of the panel regression to evaluate the relationship between private and government consumption.

The result of the CUSUM test is presented in table 3. The estimated CUSUM statistic is shown in panel A, while critical values are shown in panel B. The null hypothesis of the test is that all panel members are cointegrated, which is tested against the alternative hypothesis of non-cointegration in a subset of the panel members. The CUSUM test allows for heterogenous cointegration vectors among the panel members and, therefore, equation (28) is run separately for each country. Consistently estimated residuals are achieved by applying the DOLS estimator in the regressions.¹⁷ Applying the cointegration test to the OECD data yields a test statistic of 3.01. Comparing this statistic to the analytical critical values in panel B leads to the rejection of the null hypothesis. However, in section 5, we found that a bootstrap of the test resulted in critical values substantially larger than the analytical ones. A bootstrap of the test is also appropriate since it is able to model cross-sectional dependence in the data. Using the bootstrapped critical values, we do not reject the null hypothesis of cointegration at any conventional significance level. Thus, we find support for the hypothesis of a cointegration relation between private consumption, government consumption and GDP in the OECD countries. With this result in hand, we proceed and analyse the estimated parameters from the panel cointegration regression.

To evaluate the relationship between private and government consumption, the next step in the analysis is to see if we can find a single relation between the variables that is valid for all countries, or if we have to analyse each country separately. The relation between

¹⁷All the results presented in this section are from estimations using the DOLS estimator, setting the number of lags and leads to 2. Changing the number of lags and leads does not change the results. We have also estimated the model using the FMOLS estimator and all results are valid with this estimation method as well. We choose to present the DOLS estimates in the light of the results in Kao and Chiang (2000), who find that DOLS is more preferable in finite samples.

Table 3: The result of the CUSUM test

Panel A: The CUSUM test			
CUSUM	3.01		
Panel B: Critical values			
	10%	5%	1%
Analytical	1.28	1.64	2.33
Bootstrapped	3.45	4.04	5.46

Note: The analytical critical values are from the standard normal distribution. The bootstrapped critical values in the table are for a non-block bootstrap. For details, see section 5.

private and government consumption depends on a variety of circumstances, such as the size of the government sector and the composition of government consumption. Since these circumstances vary between countries, it is not certain that estimating a single relationship for all countries is appropriate. An aggregate relation between private and government consumption can be found by pooling the data and estimate equation (28). The result of the estimation of this aggregated relation is found in panel A in table 4. However, if the relation between private and government consumption differs among countries, it may not be adequate to estimate only one relation. We address this issue by applying a Wald test to the panel estimation of equation (28) to test if the parameters in the cointegration relation, i.e. α , β_1 and β_2 , are equal across all countries. The results of these Wald tests are presented in panel B of table 4.

For the pooled estimation, the parameter measuring the relation between private and government consumption, β_1 , is estimated at -0.59 , which is significant at the 5 percent level. The negative coefficient implies that government consumption is a substitute for private consumption. The result is very similar to the result in Ho (2001), who estimates the relation to be -0.54 for 24 OECD countries between 1981 and 1997.¹⁸ According to the Wald tests in panel B, however, the application of a single cointegration relation is dubious.

¹⁸The panel in Ho (2001) consists of 24 OECD countries. Compared to our study, Argentina and Turkey are included in Ho's analysis, while Australia is not.

Table 4: The results of a pooled estimation and Wald tests of equal coefficients

Panel A: Pooled estimation				
	α	β_1	β_2	
	1637.02	-0.59	0.62	
	(11.11)	(-12.31)	(48.88)	

Panel B: Wald test of equal coefficients				
	$\alpha_i = \alpha$	$\beta_{1i} = \beta_1$	$\beta_{2i} = \beta_2$	$\alpha_i = \alpha$ $\beta_{1i} = \beta_1$ $\beta_{2i} = \beta_2$
Wald statistic	60264.55	1148.71	1788.97	1149.34
p-value	0.00	0.00	0.00	0.00

Note: In panel A, t-values are in parentheses. The p-values in panel B are from the χ^2 distribution.

Four Wald tests are conducted. First, we test separately if α_i , β_{1i} and β_{2i} are equal to the common coefficients α , β_1 and β_2 , respectively. Second, we test the joint hypothesis that $\alpha_i = \alpha$, $\beta_{1i} = \beta_1$ and $\beta_{2i} = \beta_2$. As seen in panel B, we reject all four null hypotheses in favour of the alternative hypotheses that the coefficients are not equal for the countries. Hence, we have no reason to believe that there exists one common cointegration relation in the OECD countries and we continue the analysis by studying each country separately.

The results of separate estimations of the cointegration relation for each country are presented in table 5. Generally, there are substantial differences among the parameter estimates between the countries. The β_2 -parameter is positive and significant at the 5 percent level for all countries but one. The positive sign of this parameter is in line with our expectations, since a higher income is expected to lead to an increase in private consumption. We would expect the estimates to be somewhat larger than in the table, but the relatively low estimates are probably due to the fact that we have included GDP as the measure of income in the analysis. Since GDP is a measure of the total income in the economy, which also includes taxes, this affects the estimations downwards. If we were able to use disposable income instead of GDP, the parameters reflecting the relation between private consumption and income would probably be larger. The one country with

a negative β_2 -parameter is Luxembourg, where β_2 is estimated at -0.22 . The parameter is significant at the 5 percent level. An explanation for this strange result is probably that private consumption in Luxembourg largely follows gross national income, rather than gross domestic product. The correlation between GNI and GDP is much lower for Luxembourg than the rest of the countries with the possible exception of Ireland.

Turning to the β_1 -parameter; it is estimated to be positive for most countries. A positive parameter is found in 15 of the 23 countries and, among those, 8 estimates are significant at the 5 percent level. In 8 countries, β_1 is negative, but only 2 of these negative coefficients are significant. The general picture that arises is that private and government consumption more often can be regarded as complements rather than substitutes. In Iceland and Portugal, we find a significant relation characterized by substitutability, but these countries are exceptions. The country-by-country outcome also contradicts the results of the earlier pooled panel regression, where the aggregate relation between the variables indicates substitutability. Again, this highlights the need to be careful when analysing a panel of countries and not assume that there exists one one relation among them without testing.

One aspect that also needs to be considered in the estimations is the presence of structural breaks in the data. The re-unification of Germany is an obvious example. We have tried a specification where a dummy variable for Germany, taking the value of zero before 1990 and one afterwards, has been included in the model. The inclusion of such a dummy variable does not change the outcomes of either the CUSUM test or the estimations of the β_1 - and β_2 -parameters for Germany.

The above results support the results in Karras (1994), who also found that private and government spending can be regarded as complements for most countries. He studies another group of countries and sample period and uses another estimation method. For the countries included in both studies, the magnitude of the complementary effects are generally smaller in our study than in Karras'. A complementarity relation between the variables is also found by Amano and Wirjanto (1994) in a cointegration model of Canada. The reverse is found in, for example, Aschauer (1985) and Ahmed (1986) and Katsaitis

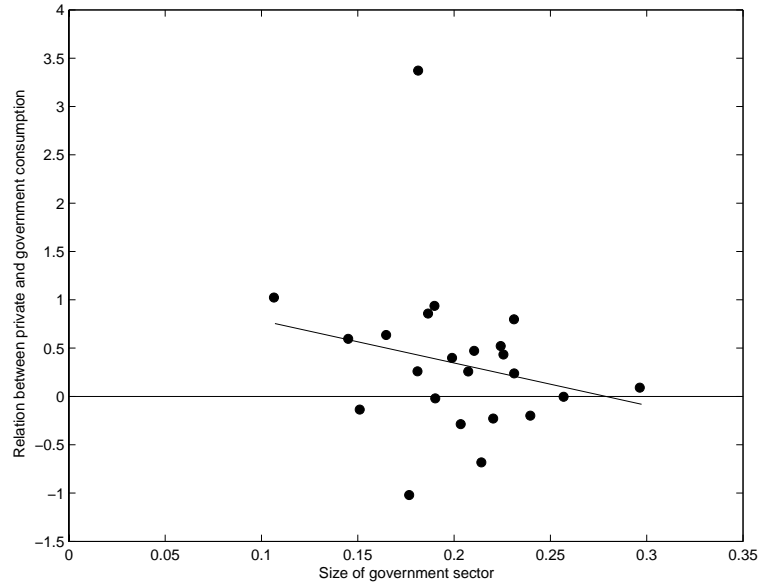


Figure 1: The relation between β_1 and the size of the government sector

(1987), who find that government spending is a substitute for private spending.

In figure 1, the results of the estimations is presented in a different way. The estimates of β_1 is related to the size of the government sector in each country. The size of the government sector is measured as an average of the ratio of government consumption to GDP between 1970 and 2001. In the figure, we see that there is a negative correlation between the estimate of β_1 and the average size of the government sector. A linear trend line is included in the graph to illustrate this negative correlation, which implies that for countries with a larger government sector, government consumption becomes more of a substitute for private consumption.¹⁹ The negative relation is in line with economic intuition and is also found by Karras (1994), who studies the relation between the size of the government sector and the degree of substitutability or complementarity between private and government consumption. The correlation between β_1 and the size of the government sector is probably due to the fact that different components of government consumption are expected to have different effects on private consumption. Some components of government consumption,

¹⁹The outlier observation in the top of the graph is Luxembourg. The negative correlation does not hinge upon this observation, nor on any other single observation.

Table 5: Parameter estimates for the OECD countries

	α	β_1	β_2
Australia	870.91 (3.64)	0.86 (4.00)	0.40 (9.07)
Austria	-1058.34 (-7.76)	0.47 (3.63)	0.51 (22.44)
Belgium	-921.46 (-5.14)	0.80 (6.96)	0.42 (20.19)
Canada	535.92 (-1.25)	0.52 (3.43)	0.49 (24.63)
Switzerland	3071.30 (1.95)	1.02 (1.76)	0.38 (3.18)
Germany	-1761.97 (-4.21)	0.94 (2.51)	0.47 (5.88)
Denmark	2254.95 (9.62)	-0.01 (-0.02)	0.41 (12.10)
Spain	725.77 (2.93)	-0.14 (-0.95)	0.59 (14.96)
Finland	890.88 (4.70)	0.43 (3.21)	0.38 (11.46)
France	1632.21 (4.79)	0.24 (0.88)	0.43 (5.48)
United Kingdom	-2230.25 (-2.67)	-0.23 (-0.65)	0.79 (21.06)
Greece	-4977.62 (-5.27)	0.64 (0.98)	0.92 (6.93)
Ireland	2419.60 (6.03)	0.26 (0.98)	0.34 (10.33)
Iceland	-871.77 (-1.71)	-0.68 (-3.34)	0.76 (11.71)
Italy	-1762.30 (-7.57)	0.40 (1.84)	0.59 (14.71)
Japan	269.86 (2.14)	0.60 (3.33)	0.46 (15.83)
Luxembourg	2739.76 (5.97)	3.37 (6.60)	-0.22 (-2.70)
Netherlands	2374.24 (7.36)	-0.20 (-0.81)	0.44 (7.00)
Norway	2962.84 (14.74)	-0.29 (-1.09)	0.41 (7.49)
New Zealand	1508.31 (3.49)	-0.02 (-0.06)	0.53 (9.07)
Portugal	30.86 (0.07)	-1.02 (-2.47)	0.83 (7.27)
Sweden	3152.88 (7.78)	0.09 (0.59)	0.34 (7.69)
United States	-1715.71 (-4.55)	0.26 (1.89)	0.69 (59.66)

Note: The parameters have been estimated with the DOLS estimator. t-values are in parentheses.

especially "core" activities for a government such as the judicial system and defence, but also the infrastructure, can be thought of as being complementary to private consumption. Many other components of public spending, on the other hand, are certainly substitutes for private consumption, such as subsidized school lunches and public transports. In between these areas, many activities can have features of both complementarity and substitutability. Such examples may be health care and education. The larger the government sector, the more probable it is that the activities in the government are substitutes for private consumption. A small government sector is likely to consist mostly of complementary activities.

7 Conclusions

In this study, we analyse the relation between private and government consumption in 23 OECD countries between 1970 and 2001, a relation that has important implications for the design and analysis of fiscal policy. Specifically, we address the issue of whether private and government consumption are substitutes or complements, i.e. whether the marginal utility of private consumption increases or decreases with government consumption. This question has been analysed in a number of studies over the years, but the empirical results are mixed. Older studies often find a substitutability relation between the variables, but in recent studies, which take into account non-stationarity and cointegration in the data, government consumption has been found to be a complement to private consumption.

We apply panel cointegration analysis to investigate the relation. A panel approach is suitable since panel cointegration tests have been shown to have better statistical properties than single-equation time series tests. In the empirical application, we use the newly developed CUSUM cointegration test by Westerlund (2005). One assumption behind the test is that the individual panel members are uncorrelated, which is a restrictive assumption in empirical applications. A bootstrap of the test is therefore conducted, allowing us to model cross-sectional dependence in the data.

The result of the bootstrap is that the bootstrapped critical values are larger than the analytical ones and that the differences are substantial. Although several factors might

contribute to this, cross-sectional dependence in the panel is the most probable one. One conclusion from this study is that it is advisable to apply a bootstrap technique when using the test in applications where cross-sectional correlation is present.

Three main conclusions can be made from the empirical analysis. The first is that the relation between private and government consumption differs to a large extent among the countries. We formally test if the parameters are the same across countries, and the result of the test implies that they are not. This highlights the fact that one has to be very careful when treating a group of countries as a common unit in panel analyses. The second conclusion is that government consumption is a complement to private consumption in the majority of the OECD countries. Only for a few of the countries do we find a relation of substitutability. No statistically significant relation can be found for many of the countries. Finally, we find a correlation between the size of the government sector and the relation between private and government consumption. The larger the government sector, the more probable it is that private and government consumption are substitutes. This correlation is economically intuitive since a small government conducts activities that are mostly complements to private consumption, while a larger government engages in an increasing number of activities that are substitutes.

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Appendix

Individual ADF tests

Table 6: Individual ADF tests

	<i>pc</i>	Δpc	<i>gc</i>	Δgc	<i>gdp</i>	Δgdp
Australia	0.85	-3.95*	-0.95	-4.90*	1.44	-3.98*
Austria	-0.22	-5.40*	-2.26	-3.83*	-0.29	-4.89*
Belgium	-1.35	-4.73*	-2.20	-3.92*	-0.37	-4.88*
Canada	-1.00	-3.14*	-2.41	-2.91*	-0.17	-3.55*
Switzerland	-0.87	-4.17*	-0.82	-3.68*	-0.58	4.13*
Germany	-1.2	-4.71*	-3.13*	-5.29*	-1.37	-4.73*
Denmark	-0.56	-4.89*	-0.90	-3.89*	0.68	-4.54*
Spain	-0.14	-2.57	0.87	-3.12*	0.33	-2.57
Finland	-1.12	-2.96*	-1.90	-2.66	-1.10	-2.80
France	-0.61	-3.87*	-0.36	-4.23*	-0.27	-4.11*
United Kingdom	1.20	-2.93*	-1.35	-3.95*	0.68	-3.17*
Greece	-0.46	-3.96*	-3.02*	-2.82	-0.67	-5.20*
Ireland	1.55	-2.79	1.54	-2.22	1.69	-1.88
Iceland	-1.17	-4.00*	-1.75	-4.89*	-0.63	-3.34*
Italy	-0.47	-4.12*	-1.09	-2.88*	-0.61	-6.15*
Japan	-1.33	-3.92*	1.89	-4.35*	-1.35	-3.66*
Luxembourg	0.67	-4.39*	1.62	-3.09*	0.06	-3.40*
Netherlands	-0.58	-2.45	1.07	-4.02*	1.03	-3.29*
Norway	0.04	-3.99*	0.28	-4.34*	0.70	-4.30*
New Zealand	-0.07	-4.03*	-1.07	-7.10*	0.16	-4.78*
Portugal	0.16	-3.02*	0.40	-4.43*	0.50	-3.75*
Sweden	-0.96	-3.36*	-2.38	-3.58*	-0.31	-3.50*
United States	0.98	-3.84*	-0.72	-3.19*	-0.08	-4.11*

Note: An asterisk (*) indicates that the test is rejected at the 5 percent significance level. The 5 percent critical value for the ADF test is -2.88.