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INTRATEMPORAL SUBSTITUTION AND GOVERNMENT SPENDING: UNIT ROOT AND COINTEGRATION TESTS IN A CROSS SECTION CORRELATED PANEL

MONICA AUTERI E MAURO COSTANTINI

Dipartimento di Teoria Economica e Metodi Quantitativi Facoltà di Scienze Politiche Università di Roma "La Sapienza" e Università di Roma "La Sapienza" Facoltà di Economia

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Abstract: There is substantial empirical literature which examines the relationship between private and public consumption. The conclusions from this literature, however, are generally mixed. In this paper, we attempt to provide some additional evidence on this relationship and to augment the empirical literature by extending this issue to panel data. The empirical framework applies the panel cointegration approach under cross-sectional dependence. To estimate the intraperiod preference parameter Bai and Kao (2004) and Westerlund panel estimators (2005b) are used. Evidence from 15 European countries indicates a significant degree of substitutability between government spending and private consumption. Therefore the existence of crowding out may render the Keynesian plea for expansionary fiscal policy unconvincing.

JEL classification: C22, E21
Keywords: permanent income; government expenditure; consumption, panel unit root, cointegration

1. INTRODUCTION

The European Union (EU) is a new framework not only for monetary policy but also for the fiscal policies of its member states. Fiscal Policy remains a national competence for EU member states, but under several constraints with the aim of controlling the fluctuation of the ratio of government spending to GDP. However since the 1970s the GDP has not been fluctuating around some constant ratio, as implied by stabilization policies, but, instead, it has steeply increased. In most cases, this increase lasted until the early 1990s, when the European Monetary System (EMS) crisis and the (European Monetary Union) EMU-entry criteria brought about increased costs of debt financing and thus the need for higher fiscal discipline. A particularly relevant issue, from a policy point of view, became then the availability of conclusive evidence on the response of economic aggregates to changes in fiscal policies. Over the past five years, there has been a renewed interest in the effects of government spending on private consumption since this relationship is central to the effectiveness of fiscal policy. The multiplier process causes an increase in government spending, or any other exogenous increase in spending, to have a greater ultimate effect on the nominal level of income through price increases, real income increases, or both, depending on where the economy is, relative to full employment. This observation makes expansionary fiscal policy attractive to those who believe in government intervention in economy control. Following however the basic prediction of the Real Business Cycle model (RBC) government spending crowds out private consumption. This suggests that standard RBC models are not appropriate to examine the macroeconomic implications of fiscal policy shocks.

A large literature has been developed and the relationship between government spending and private consumption has been estimated but since the evidence is not conclusive[1] another investigation appears warranted. To this end we use a model of permanent income based on Ogaki's (1992) theoretical model, that allows for random preference shocks in the consumer's intraperiod utility

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1 See, e.g., Aiyagari 1992 , Christiano and Eichenbaum 1992, Baxter and King 1993 , Correia 1995 , Devereux 1996 , and Kollintzas and Vassilatos 2000

function and does not require the assumption of additive separability in private and public consumption.

This two-goods permanent-income model allows us to estimate the intraperiod elasticity of substitution for private and public consumption. Using data from 1970 to 2003 for 15 European countries, we extend previous analyses from times series to panel data. New cointegration panel methodologies are applied to an intraperiod first-order condition of the model and the intratemporal elasticity of substitution between private and public consumption is estimated with panel estimators that account for cross-sectional dependence (Bai and Kao, 2004; Westerlund, 2005b). Cross-sectional dependence is considered since co-movements of economies are often observed and aggregate price ratio may tend to be correlated across countries when they are driven by common disturbances (i.e monetary shocks in the Euro area may affect all European countries) .

This article is organized as follows. Section II offers a brief literature review. Section III presents a two-goods permanent-income model with an intraperiod utility function that is not separable in private and public consumption, and that allows for random taste shocks. The intratemporal cointegration relationship implied by the model is then derived and discussed. Section IV describes the econometric methodology. The results of the unit-root and cointegration tests are presented in section V as well as the outcome of the empirical estimation of the structural parameters. Section VI concludes.

2. Brief Review of the Theoretical Literature

There is substantial empirical literature which examines the relationship between private and public consumption. Bailey (1971) and Barro (1981) argue that a general model of consumption should allow for the direct effect of government purchases of goods and services on a consumer's utility. They indicate a crowding –out effect of government spending on private consumption. With the use of a permanent income model, he estimates the intertemporal elasticities of substitution and found that government expenditure tends to crowd-out private consumption. These results confirm the basic prediction of the real business cycle (RBC) model that government spending crowds out private consumption. Intuitively, an increase in government spending creates a negative wealth effect by lowering the households' permanent income. To prevent a large drop in consumption, households increase their labor supply, but this substitution effect is typically not strong enough to offset the wealth effect. As a result, consumption decreases.

Using a general model of consumption, Kormendi (1983) and Aschauer (1985) estimate a substantial degree of substitutability between private and public consumption for the United States. Ahmed (1986) corroborates this finding for the United Kingdom. Campbell and Mankiw (1990) however do not find any significant effect in a post-war data set for the US, and Karras (1994) finds complementarity between public and private consumption in a number of countries.

The question of whether private consumption and public spending are complements or substitutes has been further studied recently, by several other authors, such as Amano and Wirjanto (1997, 1998), and Okubo (2003). These studies use a partial-equilibrium approach based on Euler equations to estimate the degree of complementarity between private consumption and government spending. Overall, empirical results yield mixed evidence of complementarity. The uncertainty of results is confirmed by Ni (1985). He shows that the relationship between private and government consumption is sensitive to the choice of the utility function and the interest rate measurement.

Recently, new empirical studies (Fatàs and Mihov (2001), Mountford, A. and H. Uhlig. (2002), Perotti, R. (2002), Galì, J., J.D. Lòpez-Salido, and J. Vallés.(2004)) find that an increase in public spending leads to a significant and persistent increase in private consumption. Starting from the assumption of price flexibility, Linnemann and Schabert (2003) and Galì, J., J.D. Lòpez-Salido, and J. Vallés (2004) examine the role of government spending in a New Keynesian framework. Their results show that price stickiness by itself does not overturn the crowding-out effect of public spending on private consumption. Galì, J., J.D. Lòpez-Salido, and J. Vallés (2004) succeed in generating a positive effect of government spending on consumption. Intuitively, when prices are sticky, an increase in government spending increases aggregate demand, which in turn raises the real wage. Higher current labor income stimulates household consumption and then aggregate consumption increase.

Given these wide range of results reported in literature, another investigation of the relationship between private and public consumption appears warranted.

Starting from a permanent income model and specifically on the basis of Ogaki's (1992) theoretical model, that allows for random preference shocks in the consumer's intraperiod utility function and does not require the assumption of additive separability in private and public consumption, we estimate the relationship between public and private consumption.

This paper attempts to contribute to the literature re-examininge the relationship between private and public consumption, and providing additional empirical evidence. Previous analyses are extended from times series to panel data spanning from 1970 to 2003, and accounting 15 European countries. The selection of these countries among the EU—Austria, Belgium, Denmark, Finland, France, Germany, Greece, Ireland, Italy, Luxembourg, Netherlands, Portugal, Spain, Sweden, and the United Kingdom —was made on the basis of data availability alone.

Recently developed theories of cointegration together with new econometric methodology of estimation are applied to exploit the restriction imposed by the intraperiod first-order condition of the model (see Ogaki (1992).

By estimating a static first-order condition of the model, we measure the intraperiod elasticity of substitution between private and public EU consumption and then we re-examine whether a certain degree of substitutability exists between government spending and private consumption. In other words we examine if a crowding-out effect exists.

3. The theoretical model

This section develops the theoretical framework of the empirical analysis following the Ogaki (1992) model. According to this model, that has found wide application, e.g. Amano and Wirjanto (1996, 1997) and Chiu (2001), a representative agent maximizes the expected lifetime consumption utility function expressed as:

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

where C_t is real private consumption at time *t,* and G_t is real government expenditures in period *t.* Equation (1) is subject to a lifetime budget constraint in a complete market at period *t*. E_t is the expectations operator based on period *t* information, and $0 \le \beta \le 1$ is a discount factor. Consider the addilog utility function:

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

where α and ν are curvature parameters with α and $\nu \ge 0$, K is a scaling factor, and Λ_{C_t} and Λ_{G_t} represent random preference shocks associated with private and public consumption, respectively. By allowing for these shocks, we avoid Garber and King's (1983) assertion that the presence of random preference shocks can often yield misleading results. Hence the representative consumer maximizes the intra-period utility of (2) subject to the intratemporal budget constraint

$$P_{c,t}C_t + P_{g,t}G_t = M_t \tag{3}$$

where $P_{c,t}$ and $P_{g,t}$ are the prices of private consumption and public consumption at time *t* and M_t is the total consumption expenditure at time *t*.

In addition, we assume that the sequences of random preference shocks, Λ_{C_t} and Λ_{G_t}, are stationary or I(0) processes. An intratemporal (or static) first-order necessary condition of the above-mentioned problem states that the relative purchase price of government to private consumption is equal to the marginal rate of substitution based on the purchase of the two types of goods; i.e.,

$$\frac{C_t^{-\alpha}\Lambda_{C_t}}{P_{c,t}} = \lambda_t \tag{4}$$

$$\frac{KG_t^{-\nu}\Lambda_{G_t}}{P_{g,t}} = \lambda_t \tag{5}$$

then, combining (4) and (5), we obtain

$$\frac{P_{g,t}}{P_{c,t}} = \frac{KG_t^{-\nu}}{C_t^{-\alpha}}\frac{\Lambda_{G_t}}{\Lambda_{C_t}} \tag{6}$$

Let $P_t = P_{g,t}/P_{c,t}$, then we have:

$$P_t = \frac{KG_t^{-\nu}}{C_t^{-\alpha}}\frac{\Lambda_{G_t}}{\Lambda_{C_t}} \tag{7}$$

Taking logarithms on both sides of (7), yields the following cointegrating equation:

$$\ln P_t + \nu \ln G_t - \alpha \ln C_t = \xi_t \tag{8}$$

In (8), ξ_t collects the remaining terms and represents a stationary process of preference shocks with zero mean; therefore, (8) implies that $\ln P_t, \ln G_t \ \ln C_t$ are cointegrated with cointegrating vector $[1, \nu, \alpha]$ Rearranging (8), we obtain extra information for intratemporal elasticity of substitution between government and private goods. For example, (8) can be rewritten as:

$$\ln G_t = \left(\tfrac{\alpha}{\nu}\right) C_t - \left(\tfrac{1}{\nu}\right)\ln P_t + \varepsilon_t \tag{9}$$

The estimate ν/α is defined as the intratemporal elasticity of substitution between government and private spending. In addition recall that both G and C are treated as choice variables for a representative consumer, so that the representative consumer maximizes his (or her) utility by optimally consuming both private and government goods; then in the above mentioned intratemporal utility function $\left(\tfrac{1}{\alpha}\right)$, $\left(\tfrac{1}{\nu}\right)$ may be interpreted as the intratemporal elasticity of substitution for private and government consumption, respectively. Hence our empirical analysis focuses on estimating parameter ν and α to calculate these elasticities within a panel data set. Extending the model from time series to panel data imply the equation that we are going to estimate becomes:

$$\ln G_{i,t} = \left(\tfrac{\alpha}{\nu}\right) C_{i,t} - \left(\tfrac{1}{\nu}\right) \ln P_{i,t} + \varepsilon_{i,t} \qquad (10)$$

where $\varepsilon_{i,t}$ is a stationary process with zero mean. Ogaki and Park (1992) show that this cointegration approach allows for non-orthogonal but stationary multiplicative measurement error, the presence of liquidity constraints, and a general form of time separability in preferences. The latter result, however, holds only under the restrictive assumption of additive separability between the two goods. In the case of non-separability such as ours, the cointegration approach is not robust to time non-separability except in very few special cases. One such case is given by the intraperiod utility function of the form $U(C_t, G_t) = \Lambda_t \left(C_t^*\right)^{\alpha} \left(G_t^*\right)^{\nu}$, where $\Lambda_t = \left(\Lambda_{C_t}, \Lambda_{G_t}\right)$ and C_t^* and G_t^* are service flows from purchases of private and public goods, respectively.

4. Motivation for unit root, cointegration and estimation in panel data.

Previous analyses are extended from times series to nonstationary panel data. In finite sample, unit root test procedures are known to have limited power against the alternative hypothesis with highly persistent deviation from equilibrium (Frankel and Rose, 1996) and this problem seems to be particularly severe for small samples (Campbell and Perron, 1991). It is well known that, in small sample sizes, unit root tests generally have low power to distinguish nonstationary series from stationary series that are persistent (MacDonald, 1996). It is noteworthy that the use of panel data increases the power of the unit root tests by increasing the number of cross-sectional units (Banerjee, 1999), then in our empirical analysis, panel unit root tests for cross-sectional dependence are therefore considered[2].

Our estimation approach proceeds in three steps. In the first step we investigate the panel properties of the variables. In the second step of our empirical analysis, we apply recently developed theories of cointegration to an intraperiod

2 Panel unit root tests with cross-sectional dependencies have been widely used in international finance applications where the time dimension (t) is larger than the macroeconomic dimension (T). For macroeconomics application see Hurlin (2004).

first-order condition of the model and then, in the last step, we estimate the intratemporal elasticity of substitution.

4.1. Panel unit root

We begin by examining the panel properties of the data. To this end we first use the test developed by Choi (2004). We then apply the tests proposed by Bai and Ng (2003), Moon and Perron (2004), and then two tests provided by Pesaran (2005).

Choi proposed new panel unit root tests for cross-sectionally correlated panels. The cross-sectional correlation is modeled by error-component models. The test statistics are derived from combining p-values from the Augmented Dickey-Fuller test applied to each time series whose non-stochastic trend components and cross-correlation are eliminated by Elliot, Rothenberg and Stock's (1996) GLS-based de-trending and the conventional cross-sectional demeaning panel data. The panel unit root tests proposed are:

$$P_m = -\frac{1}{\sqrt{N}}\sum_{i=1}^{N}\left(\ln(p_i)+1\right)$$

$$Z = \frac{1}{\sqrt{N}}\sum_{i=1}^{N}\Phi^{-1}(p_i)$$

$$L* = \frac{1}{\sqrt{\pi^2 N/3}}\sum_{i=1}^{N}\ln\left(\frac{p_i}{1-p_i}\right),$$

where the P_m test is a modification of Fisher (1932) inverse chi square, $\Phi(\cdot)$ is the standard normal cumulative distribution function and p_i indicates the asymptotic p-value of one of the Dickey-Fuller-GLS test for country *i*.[3] For $T \to \infty$ and $N \to \infty$,

$$P_m, Z, L* \Rightarrow N(0,1)$$

Bai and Ng consider the following factor model:

$$Y_{it} = D_{it} + \lambda_i^{'} F_t + e_{it}$$

3 The percentiles of the asymptotic p-values of the Dickey-Fuller-GLS tests are simulated by Choi

where D_{it} is a polynomial trend function, F_{it} is an $r \times 1$ vector of common factors, and λ_t is a vector of factor loading. Thus the series Y_{it} is decomposed into three components: a deterministic one, a common component with factor structure and an idiosyncratic error component. The process Y_{it} may be non-stationary if one or more of the common factors are non-stationary, or the idiosyncratic error is non-stationary, or both. To test the stationarity of the idiosyncratic component, Bai and Ng propose to pool the individual ADF t-statistics with de-factored estimated component e_{it} in a model with no deterministic trend:

$$\Delta e_{it} = \delta_{i,0}\hat{e}_{i,t-1} + \sum_{j=1}^{p} \delta_{i,j}\Delta\hat{e}_{i,t-j} + u_{i,t}$$

Let $ADF_{\hat{e}}^{c}(i)$ be the ADF t-statistic for the *i*-th region. The asymptotic distribution of the $ADF_{\hat{e}}^{c}(i)$ coincides with the Dickey-Fuller distribution for the case of no constant. However, these individual time series tests have the same low power as those based on the initial series. Bai and Ng aimed at testing the common factor and the idiosyncratic error separately. They proposed pooled tests based on Fisher's type statistics defined as in Choi (2001) and Maddala and Wu (1999). Let $P_{\hat{e}}^{c}$ be the p-value associated with $ADF_{\hat{e}}^{c}(i)$, then

$$Z_{\hat{e}}^{c} = \frac{-2\sum_{i=1}^{N} \log p_{\hat{e}}^{c}(i) - 2N}{\sqrt{4N}} \xrightarrow{D} N(0,1)$$

Moon and Perron (MP, hereafter) developed several unit root tests in which the cross-sectional units are correlated. To model cross-sectional dependence, an approximate linear dynamic factor model is provided. The panel data are generated by both idiosyncratic shocks and unobservable dynamic factors that are common to all individual units but each individual reacts heterogeneously. In our analysis, we apply the following MP tests:

$$t_a^{*} = \frac{\sqrt{NT}\left(\hat{\rho}_{pool}^{+} - 1\right)}{\sqrt{\frac{2\hat{\phi}_e^4}{\omega_e^4}}}$$

$$t_b^* = \sqrt{NT}\left(\hat{\rho}_{pool}^{+} - 1\right)\sqrt{\frac{1}{NT^2} tr(Y_{-1} Q_B Y_{-1}^{'})}\left(\frac{\hat{\omega}_e^2}{\phi_e^4}\right)$$

where $\hat{\rho}_{pool}^{+}$ is the bias-corrected pooled autoregressive estimated of $\hat{\rho}_{pool}^{+}$, $\hat{\omega}_e^2$ and $\hat{\phi}_e^4$ are respectively the estimates of the cross sectional average of long run variance of $\hat{e}_{it}$ and the cross sectional average of $\hat{\omega}_{e,i}^4$.[4]

To deal with the problem of cross-sectional dependence, Pesaran does not consider the deviations from the estimated common factor, but he proposed to augment the standard DF (or ADF) regression with the cross section averages of lagged levels and first-differences of the individual series. The panel unit root tests are then based on the average of individual cross-sectionally augumented ADF statistics (CADF). The individual CADF statistics may be used to construct modified version of the t-bar test developed by Im, Pesaran and Shin (2003, **IPS** hereafter), the inverse chi-square test (P test) developed by Maddala and Wu (1999) and the inverse normal test (Z test) proposed by Choi(2001). Pesaran proposes a truncated version of the test to avoid undue influences of extreme outcomes that could emerge in the case of small T. The simple average of cross-sectionally augmented IPS test and its truncated version are :

$$CIPS(N,T) = N^{-1}\sum_{i=1}^{N} t_i(N,T)$$

and

$$CIPS*(N,T) = N^{-1}\sum_{i=1}^{N} t_i^*(N,T)$$

where $t_i(N,T)$ and $t_i^*(N,T)$ are the cross-sectionally augmented Dickey-Fuller statistic for the ith cross section unit and the truncated version respectively given by the t-ratio of the coefficient of $y_{i,t-1}$ in the CADF regression:

$$\Delta y_{it} = a_i + b_i y_{i,t-1} + c_i y_{t-1} + d_i \Delta y_t + e_{it}$$

4 For details on Moon and Perron tests see the appendix B.

4.2. Panel Cointegration

The theory outlined in the previous section, together with the panel unit-root tests, imply that the variables considered are cointegrated. To examine whether consistent evidence with cointegration exists, we move to the second step, and we apply a set of panel cointegration tests. We start with the ADF panel cointegration test (Kao, 1999), which assumes that the cointegrating vectors are homogeneous.

Let $\hat{e}_{it}$ be the estimated residual from the following regression:

$$y_{it} = \alpha_i + \beta x_{it} + e_{it}$$

The ADF test is applied to the estimated residual:

$$\hat{e}_{it} = \gamma \hat{e}_{i,t-1} + \sum_{j=1}^{p} J_j \Delta \hat{e}_{i,t-j} + v_{i,tp}$$

where *p* is chosen so that the residual $v_{i,tp}$ are serially uncorrelated. The ADF test statistic is the usual t-statistic of $\rho = 1$ in the previous equation. With the null hypothesis of no cointegration, the ADF test statistics can be constructed as:

$$ADF = \frac{t_{ADF} + \left(\sqrt{6N}\,\hat{\sigma}_v / 2\,\hat{\sigma}_{0v}\right)}{\sqrt{\left(\hat{\sigma}_{0v}^2 / 2\,\hat{\sigma}_v^2\right) + \left(10\hat{\sigma}_{0v}^2\right)}}$$

where t_{ADF} is the t-statistic of γ in the ADF regression, $\hat{\sigma}_v^2 = \Sigma_{u\varepsilon} - \Sigma_{u\varepsilon}\Sigma_{\varepsilon}^{-1}$ and $\hat{\sigma}_{0v}^2 = \Omega_{u\varepsilon} - \Omega_{u\varepsilon}\Omega_{\varepsilon}^{-1}$, Ω is the long run covariance matrix and t_{ADF} is the t-statistic in the ADF regression. Kao shows that the ADF test converges to a standard normal distribution N(0,1).

After performing the ADF test, the Durbin-Hausman panel test (DH_p) proposed by Westerlund (2005a) is then applied[5]. This test allows for cross-sectional dependence that it is modeled by a factor model in which the errors of the equation (10) are generated by both idiosyncratic innovations and unobservable factors that are common across the members of the panel. In other words, in the equation (10) the errors are modeled as follows:

$$\varepsilon_{i,t} = F_T \lambda_i + u_{it}\,,$$

[5] Westerlund (2005a) also proposed a group mean test.

where F_T is a $1 \times K$ vector of common factors and λ_i is a conformable vector of factor loading. The common factor F_T plays the role of reducing the dimensionality of the cross-sectional covariance structure of $\varepsilon_{i,t}$. The extent of this dependency is determined by λ_i.[6]

To investigate the existence of the cointegrating relationship in the equation (10), we have to test whether u_{it} is I(1) or not. To this end, we first estimate the equation (10) by OLS and then we estimate the common factors by applying the principal components method to the OLS residuals. A test of no cointegration can be developed by subjecting the defactored residuals to a unit root test. This approach is valid if u_{it} is stationary. Thus, testing the null hypothesis of no cointegration is equivalent to testing whether $\rho_i = 1$ in the following autoregression:

$$\hat{u}_{it} = \rho_i \hat{u}_{it-1} + z_{it}$$

Let $E_{it} = \left(\hat{u}_{it}, \hat{u}_{it-1}, \Delta\hat{u}_{it}\right)'$, $E_i = \sum_{t=1}^{T} E_{it} E_{it}^{'}$ and $E = \sum_{t=1}^{T} E_i$. The panel test is constructed under the maintained assumption that $\rho_i = \rho$ for all *i*. The Durbin-Hausman panel test is defined as follows:

$$DH_p \equiv \hat{\sigma}^2 \hat{\gamma}_0^{-2} \left(\tilde{\rho} - \hat{\rho}\right) E_{22}$$

where $\tilde{\rho} = E_{12}^{-1} E_{11}$, $\hat{\rho} = E_{22}^{-1} E_{12}$, $\hat{\sigma}_i^2 = \sum_{k=-M}^{M} \omega\left(k/M\right) \hat{\gamma}_{ik}$, $\hat{\gamma}_{ik} = T^{-1} \sum_{t=k+1}^{T} \hat{z}_{it} \hat{z}_{it+k}$, $\hat{\sigma}^2 = N^{-1} \sum_{i=1}^{N} \hat{\sigma}_i^2$, $\hat{\gamma}_k = N^{-1} \sum_{i=1}^{N} \hat{\gamma}_{ik}$, $\hat{z}_{it} = \hat{u}_{it} - \hat{\rho}_i \hat{u}_{it-1}$ and $\omega\left(k/M\right) = 1 - k/(1+M)$.

For the DH_p test, the null and the alternative hypothesis is designed as $H_0 : \rho_i = 1$ for all *i* against $H_0 : \rho_i = \rho$ and $\left|\rho\right| < 1$ for all *i*. A rejection of the null hypothesis should therefore be taken as evidence in favour of cointegration for all individuals in the panel.

The Durbin-Hausman statistics are derived under the condition that K, the number of common factors, is known. When it is unknown, a feasible approach is to consider the estimation problem as a model selection issue and estimate K

[6] For further details on the assumption of the test see Westerlund (2005a).

by minimizing an information criterion. The criterion used in this paper is defined as follows:

$$IC_{p2}(K) = \ln(V(K)) + K\left(\frac{N+T}{NT}\right)\ln(C_{NT}^2),$$

where $V(K) = \left(NT\right)^{-1}\sum_{i=1}^{N}\sum_{t=1}^{T}\Delta u_{it}$ and $C_{NT} = \min\left\{N^{1/2}, T^{1/2}\right\}$

For the consistency of the test, the bandwidth has not to increase too fast than T. To this end, we choose M to the largest integer less than $4(T/100)^{2/9}$, as suggested by Newey and West (1994).

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4.3 Parameter Estimation

After we examine whether consistent evidence with cointegration exists, as a last step of our analysis, we estimate the structural parameters to finally calculate the intratemporal elasticity of substitution between private and public consumption with panel estimators recently proposed by Bai and Kao (2004) and Westerlund's (2005b). Westerlund studying the small-sample properties of Bai and Kao's (2004) fully modified (FM) estimator when the number of factors is unknown, proposes a bias-adjusted OLS estimator. This estimator can be employed to obtain unbiased estimates of a cointegrated panel data regression with cross-sectional dependence. As in Bai and Kao (2004), cross-correlation is modelled using a common factor structure. The number of factors are endogenously determined from the data using several new panel information criteria. Westerlund (2005b) shows that the OLS estimator is indeed severely biased when the errors are cross-sectionally correlated. However, the bias-adjusted OLS and FM estimators can provide more accurate estimates.

If we consider the following fixed effect panel regression:

$$y_{it} = \alpha_i + \beta x_{it} + e_{it}, \quad \text{i = 1, ... , n , t= 1, ... T}$$

where y_{it} is $k\times 1$, β is a $1\times k$ vector of the slope parameters, x_{it} is a $k\times 1$, integrated processes of order one for all *i* , $x_{it} = x_{it-1} + \varepsilon_{it}$, α_i is the intercept, and e_{it} is the stationary regression error that is generated by the following factor model:

$$e_{it} = \lambda_i^{'} F_t + u_{it}$$

where F_t is a $r \times 1$ vector of common factors, $\lambda_i^{'}$ is a $r \times 1$ vector of factor loading and u_{it} is the idiosyncratic component of e_{it}, which means $E(e_{it}e_{jt}) = \lambda_i^{'} E(FF_t^{'})\lambda_j$. For ε_{it}, we could also have the following factor structure $\varepsilon_{it} = \gamma_i^{'} F_t + \eta_{it}$

The OLS estimator is:

$$\hat{\beta}_{OLS} = \left[\sum_{i=1}^{n} \sum_{i=1}^{T} y_{it} \left(x_{it} - \overline{x}_i \right)^{'} \right] \left[\sum_{i=1}^{n} \sum_{i=1}^{T} \left(x_{it} - \overline{x}_i \right) \left(x_{it} - \overline{x}_i \right)^{'} \right]^{-1}$$

where $\overline{x}_i = \frac{1}{N} \sum_{i=1}^{N} x_{it}$

The bias-adjusted estimator developed by Westerlund (2005b) is:

$$\hat{\beta}^{+} = \left[\sum_{i=1}^{n} \sum_{i=1}^{T} y_{it} \left(x_{it} - \overline{x}_i \right)^{'} \right] \left[\sum_{i=1}^{n} \sum_{i=1}^{T} \left(x_{it} - \overline{x}_i \right) \left(x_{it} - \overline{x}_i \right)^{'} \right]^{-1} - B_{NT}$$

where $B_{NT} = \frac{6}{T} \left(\frac{1}{N} \sum_{i=1}^{N} \lambda_i^{'} \left(-\frac{1}{2} \Omega_{i23} + \Delta_{i23} \right) - \frac{1}{2} \Omega_{13} + \Delta_{13} \right) \Omega_{33}^{-1}$ and $\overline{x}_i = \frac{1}{N} \sum_{i=1}^{N} x_{it}$.

Since Ω_i, Δ_i and λ_i are unobservable, $\hat{\beta}^{+}$ cannot be estimated. When the true number of factors are known, a feasible version of the bias-adjusted OLS estimator can be obtained. In a first step, F_t and λ_i are estimated by the principal component method. Then the estimates of F_t, λ_i and e_{it} are used to construct consistent estimators $\hat{\Omega}_i$ and $\hat{\Delta}_i$ of Ω_i and Δ_i. These are then used to get feasible bias-adjusted OLS estimator.

Bai and Kao, 2004 pointed out that an iterative procedure may lead to gain efficiency in estimating β. Westerlund proposes the following recursive estimation. First β is estimated using OLS. The estimates of β are then used to obtain $\hat{\lambda}_i$, $\hat{\Omega}_i$, $\hat{\Delta}_i$ and subsequently $\hat{\beta}^{+}$. The next step is to re-estimate $\hat{\lambda}_i$, $\hat{\Omega}_i$, $\hat{\Delta}_i$ based on $\hat{\beta}^{+}$. These estimates are then used to update $\hat{\beta}^{+}$. Then, the iteration process is conducted until convergence is reached or until the number of iterations reaches some predetermined upper boundary.

The last estimator we used is the FM estimator that is the OLS estimator $\hat{\beta}_{OLS}$. Corrected for endogeneity and serial correlation to The endogeneity correction is achieved by modifying the variable y_{it} with the following transformation:

$$y_{it}^{+} = y_{it} - \left(\lambda_i^{'}\Omega_{F\varepsilon i} + \Omega_{u\varepsilon i}\right)\Omega_{\varepsilon i}^{-1}\Delta x_{it},$$

where Ω is the long run-variance matrix of $\omega_{it} = \left(F_t^{'}, u_{it}, \varepsilon_{it}^{'}\right)'$

The serial correlation correction term has the form:

$$\Delta_{b\varepsilon i}^{+} = \Delta_{b\varepsilon i} - \Omega_{b\varepsilon i}\Omega_{\varepsilon i}^{-1}\Delta_{\varepsilon i}$$

the infeasible FM estimator is:

$$\tilde{\beta}_{FM} = \left[\sum_{i=1}^{n}\left(\sum_{i=1}^{T} y_{it}^{+}\left(x_{it} - \overline{x}_i\right)^{'} - T\left(\lambda_i^{'}\Delta_{F\varepsilon i}^{+} + \Delta_{u\varepsilon i}^{+}\right)\right)\right]\left[\sum_{i=1}^{n}\sum_{i=1}^{T}\left(x_{it} - \overline{x}_i\right)\left(x_{it} - \overline{x}_i\right)^{'}\right]^{-1}$$

The feasible FM estimator, $\hat{\beta}_{FM}$, is obtained by substituting λ, F, $\sum_i$ and Ω_i with $\hat{\lambda}$, $\hat{F}$, $\hat{\sum}_i$ and $\hat{\Omega}_i$:

$$\hat{\beta}_{FM} = \left[\sum_{i=1}^{n}\left(\sum_{i=1}^{T} \hat{y}_{it}^{+}\left(x_{it} - \overline{x}_i\right)^{'} - T\left(\hat{\lambda}_i^{'}\hat{\Delta}_{F\varepsilon i}^{+} + \hat{\Delta}_{u\varepsilon i}^{+}\right)\right)\right]\left[\sum_{i=1}^{n}\sum_{i=1}^{T}\left(x_{it} - \overline{x}_i\right)\left(x_{it} - \overline{x}_i\right)^{'}\right]^{-1}$$

where $\hat{y}_{it}^{+} = y_{it} - \left(\hat{\lambda}_i^{'}\hat{\Omega}_{F\varepsilon i} + \hat{\Omega}_{u\varepsilon i}\right)\hat{\Omega}_{\varepsilon i}^{-1}\Delta x_{it}$,

The two-step FM estimator (2S-FM) is obtained by using $\hat{\Omega}^{(1)}$ and $\hat{\lambda}^{(1)}$:

$$\hat{\beta}_{2S}^{1} = \left[\sum_{i=1}^{n}\left(\sum_{i=1}^{T} \hat{y}_{it}^{+(1)}\left(x_{it} - \overline{x}_i\right)^{'} - T\left(\hat{\lambda}_i^{'(1)}\hat{\Delta}_{F\varepsilon i}^{+(1)} + \hat{\Delta}_{u\varepsilon i}^{+(1)}\right)\right)\right]\left[\sum_{i=1}^{n}\sum_{i=1}^{T}\left(x_{it} - \overline{x}_i\right)\left(x_{it} - \overline{x}_i\right)^{'}\right]^{-1}.$$

The iterative FM estimator (CUP-FM) is obtained by estimating parameters and long-run covariance matrix and loading recursively. Therefore, $\hat{\beta}_{FM}$, $\hat{\Omega}$ and $\hat{\lambda}$ are estimated repeatedly, until convergence is reached:

$$\hat{\beta}_{CUP} = \left[\sum_{i=1}^{n}\left(\sum_{i=1}^{T} \hat{y}_{it}^{+}\left(\hat{\beta}_{CUP}\right)\left(x_{it} - \overline{x}_i\right)^{'} - T\left(\hat{\lambda}_i^{'}\left(\hat{\beta}_{CUP}\right)\hat{\Delta}_{F\varepsilon i}^{+}\left(\hat{\beta}_{CUP}\right) + \hat{\Delta}_{u\varepsilon i}^{+(1)}\left(\hat{\beta}_{CUP}\right)\right)\right)\right]\left[\sum_{i=1}^{n}\sum_{i=1}^{T}\left(x_{it} - \overline{x}_i\right)\left(x_{it} - \overline{x}_i\right)^{'}\right]^{-1}$$

Bai and Kao show that the CUP-FM has a superior small sample properties with the 2S-FM estimator.

5. Data and empirical results

The data set, which spans the period of 1970 to 2003, is constructed with annual data of 15 European countries[7] taken from OECD Statistical compendium (2004). The per-capita consumption series is obtained by dividing personal real consumption of nondurable goods and services by the total population of age 16 and over. The per-capita government expenditure series is measured as the ratio of national real government purchases of goods and services to the total population of age 16 and over. The relative price measure is simply the ratio of public to private expenditure prices where government and private price are respectively the deflator of government and private consumption.

We begin by examining the properties of the data. To this end, we first formally test the null hypothesis of unit root in panel data using the previously discussed approaches. The unit root test statistics are reported in Tables 1, 2 and 3. For the variable $\ln G_{i,t}$ the null hypothesis of a unit root cannot be rejected at the 5% level of significance (only the Bai and Ng tests reject the null hypothesis). The tests show strong evidence of unit root process in the data and therefore we conclude that the variables under consideration are well characterized as nonstationary or I(1) processes.

For private consumption, $\ln C_{i,t}$, the null hypothesis of nonstationarity cannot be rejected at 5 % level, except for MP and Pesaran tests (and in this last case only for 1 lag). The rejection of the null hypothesis does not imply that the nonstationarity is rejected for the idiosyncratic component of all European countries. It means that the null hypothesis is rejected for a sub-group of countries. In addition, the rejection of the nonstationarity of the idiosyncratic component does not imply that the series is stationary, since the common factor

[7] Given the sample period and the lack of reliable time series for many countries, in this paper we do not consider the countries that joined the EU on May 1st, 2004. The countries selection —Austria, Belgium Denmark, Finland, France, Germany, Greece, Ireland, Italy, Luxembourg, Nederland, Portugal, Spain, Sweden, and the United Kingdom —was made on the basis of data availability alone.

may be non-stationary. It is noteworthy that the Moon and Perron test is more radical and it does not test for the unit root in common factors.

With respect to the variable $\ln P_{i,t}$, the relative price ratio, the null hypothesis of nonstationarity can be rejected only in the case of the MP tests at 5 % significance level. In explaining the MP's findings, suggestions similar to those provided in the previous case, can be applied.

When the presence of nonstationarity is confirmed, the empirical relationship between the three variables is subject to cointegration analysis.

The theory outlined in the previous section, together with the unit-root test results, imply that the variables $\ln G_{i,t}$, $\ln C_{i,t}$, $\ln P_{i,t}$, may be cointegrated . Table 4 presents panel cointegration test results. All tests show evidence of cointegrating relationship between the three variables at 5 % level. In other words, the equilibrium errors in the equation (10) are stationary and the test results are consistent with cointegration between the three series, which would suggest evidence in favor of the model. The equation then may be estimated.

Overall the results are encouraging. In tables 5-6 two-step and iterative procedure results are reported.

It is noteworthy that the evidence from the cointegration approach suggests that intratemporal substitution is an important future of EU private and public consumption behavior.

The parameter that gains importance in this process is the within period (intratemporal) elasticity of substitution $\left(\frac{\alpha}{\nu}\right)$ between government and private spending. The larger the elasticity the greater the substitution towards private and government consumption will be. The cointegrating parameter estimates admit an intratemporal elasticity of substitution $\left(\frac{\alpha}{\nu}\right)$ that is roughly centered on 1. From the above mentioned estimates we are able to calculate in addition $\left(\frac{1}{\alpha}\right)$, $\left(\frac{1}{\nu}\right)$ that may be interpreted as the intratemporal elasticity of substitution for private and government consumption, respectively and appear to be smaller than 0.5

Then, the estimate close to one[8] seems to indicate an effective intratemporal elasticities of substitution between $\ln G_{i,t}$ and $\ln C_{i,t}$ and therefore indicates an effective fiscal multiplier on stimulating private consumption. In a nutshell, we find that the intratemporal elasticity of substitution implies that fiscal expansion is consistently effective.

The economic plausibility of this preference parameter estimate provides some further support for the model. It should be stressed that estimates based on Eq.(10) are average effects over the sample period. That is, unless the composition of both private and government spending remains stable, the permanent-income approach to government expenditures does not assume that the relationship will be stable over any particular subperiod.

Overall the results are therefore encouraging. There is, however, an important qualification to our results. We do not control the possible effects of binding liquidity constraints for some consumers. This omission may be an important factor, since a number of researchers have attributed rejections of the permanent-income model to the possible presence of liquidity constraints. Bean (1986) and Cushing (1992), for example, find empirical support for the hypothesis that a significant proportion of U.S. consumers face binding liquidity constraints. Evans and Karras (1992) use annual data from 54 countries and find liquidity constraints to be present in many of the countries.

6. Conclusion

In this paper, we attempt to provide some additional empirical evidence on the relationship between private and public consumption. We consider a two-goods permanent-income model that allows for random preference shocks and does not require the assumption of additive separability in the consumption of two goods. Employing recent cointegration panel estimators (Bai and Kao, 2004; Westerlund, 2005b) to examine this issue in panel data with dynamic panel

[8] More precisely using the OLS estimator we found an intratemporal elasticity of substitution $\left(\frac{\alpha}{\nu}\right)$ equal to 0.99, and with the other estimators we found an intratemporal elasticity of substitution even consistently greater than one.

estimators that account for cross-sectional dependence, we estimate a static first-order condition of the model using a cointegration approach we are able to measure the intraperiod elasticity of substitution between private and public EU consumption.

In addition to the substantial evidence for accepting cointegration, the present analysis first suggests that intratemporal elasticity of substitution is an important feature of European economies. The empirical results suggest that the intraperiod elasticity of substitution for two consumption goods is about 1. This indicates a degree of substitutability between government spending and private consumption, or the crowding-out effect.

The existence of the crowding-out effect would make the government spending multiplier smaller than it is anticipated. In fact the crowding-out phenomenon describes the process whereby an increase in government spending decreases other components of aggregate demand, thus reducing the government spending multiplier effect on stimulating aggregate demand. Besides the level of employment, the crowding-out effect is related to the means used to finance an increase in government spending. If taxes are used to finance an increase in government spending, this multiplier is called the balanced-budget multiplier, reflecting the fact that the fiscal action has no impact on the size of the government's budget deficit or surplus. In this case, consumers reduce consumption spending to be able to pay the higher taxes. The decrease in consumption demand partially offsets the increase in government spending, reducing the size of the multiplier. Moreover, the multiplier process usually assumes that government sells bonds to finance an increase in its spending, in this case, extra crowding out comes about. However EU countries are strictly limited in their use of the fiscal policies. In addition the Stability Pact reduce the national authorities power to use fiscal policies as a proxy for monetary policy and extensive intergovernmental transfers. The Stability and Growth Pact limits Member State borrowing up to 3 per cent of GDP unless an economy is in recession. A system of heavy fines can be imposed should the rules be broken.

The empirical results seem to provide some robust evidence to support the argument that in a general model of consumption there is a direct effect of government consumption on an agent’s utility. This implies that the crowding out effect renders the Keynesian plea for expansionary fiscal policy unconvincing.

In addition the plausibility of these estimates suggests that the Ogaki (1992) model is a useful tool for estimating parameters not only with time series but also in a panel data framework

APPENDIX

A. Bai and Ng panel unit root test

Consider the following model with individual effect and without time trend:

$$y_{it} = \alpha_i + \lambda_i^{'} F_t + e_{it} \qquad \text{(A.1)}$$

where F_t is a $r \times 1$ vector of common factors and λ_i is a vector of factor loadings.[9] Among the *r* common factors, we allow *r₀* and *r₁* to be stochastic common trends with $r_0 + r_1 = r$. The corresponding model in first difference is:

$$\Delta y_{it} = \lambda_i^{'} F_t + z_{it} \qquad \text{(A.2)}$$

where $z_{it} = \Delta e_{it}$ and $f_t = \Delta F_{it}$ with $E\left(f_t\right) = 0$. Applying the principal-components approach to Δy_{it} yields *r* estimated factors $\hat{f}_t$, the associated loadings $\hat{\lambda}_i$, and the estimated residuals, $z_{it} = y_{it} - \hat{\lambda}_i^{'} \hat{f}_t$.

If we define:

$$\hat{e}_{it} = \sum_{s=2}^{t} \hat{z}_{it}$$

$\hat{F}_t = \sum_{s=2}^{t} \hat{z}_{it}$, an $r \times 1$ vector, for t = 2,….T,

then we have:

1. Let $ADF_{\hat{e}}^{c}(i)$ be the t statistics for testing $d_{i0} = 0$ in the univariate augumented autoregression (with no deterministic terms):

$$\Delta \hat{e}_{it} = d_{i0} \hat{e}_{it-1} + d_{i1} \Delta \hat{e}_{it-1} + ... + d_{ip} \Delta \hat{e}_{it-p} + error \qquad \text{(A.3)}$$

2. If $r = 1$, let $ADF_{\hat{F}}^{c}$ be the *t* statistics for testing $\delta_{i0} = 0$ in the univariate augumented autoregression (with an intercept):

$$\Delta F_t = c + \delta_0 \hat{F}_{t-1} + \delta_{i1} \Delta \hat{F}_{t-1} + ... + \delta_p \Delta \hat{F}_{t-p} + error \qquad \text{(A.4)}$$

[9]Specifically, the idiosyncratic error follows this process: $\left(1 - \rho_i L\right) e_{it} = D_i \left(L\right) \in_{it}$.

3. If r>1, demean $\hat{F}_t$ and denote $\hat{F}_t^c = \hat{F}_t - \bar{\hat{F}}$, where $\bar{\hat{F}} = (T-1)^{-1}\sum_{t=2}^{T}\hat{F}_t$. Start with $m = r$;

   A. $\hat{\beta}_\perp$ denotes the *m* eigenvectors associated with the *m* largest eigenvalues of $T^{-2}\sum_{t=2}^{T}\hat{F}_t^c\hat{F}_t^{c'}$ and $\hat{Y}_t^c = \hat{\beta}_\perp^{'}\hat{F}_t^c$. Two different statistics may be considered:

   B.I Let $K(j) = 1 - j/(j+1)$, j = 0, 1,…….J:

   i) Let $\hat{\xi}_t^c$ be the residuals from estimating a first-order VAR in $\hat{Y}_t^c$. In addition, let

$$\hat{\Sigma}_1^c = \sum_{j=1}^{J}K(j)\left(T^{-1}\sum_{t=2}^{T}\hat{\xi}_{t-j}^c\hat{\xi}_t^{c'}\right)$$

   ii) Let $v_c^c(m)$ be the smallest eigenvalue of

$$\Phi_c^c(m) = 0.5\left[\sum_{t=2}^{T}\left(\hat{Y}_t^c\hat{Y}_{t-1}^{c'} + \hat{Y}_{t-1}^c\hat{Y}_t^{c'}\right) - T\left(\hat{\Sigma}_1^c + \hat{\Sigma}_1^{c'}\right)\right]\left(\sum_{t=2}^{T}\hat{Y}_t^c\hat{Y}_{t-1}^{c'}\right)^{-1} \quad \text{(A.5)}$$

   iii) Define $MQ_c^c(m) = T\left[\hat{v}_c^c(m) - 1\right]$.

   B.II For *p* fixed that does not depend on N and T

   i) Estimate a VAR or order p in $\Delta\hat{Y}_t^c$ to get $\hat{\Pi}(L) = I_m - \hat{\Pi}_1 L - .... - \hat{\Pi}_p L_p$ and filter $\hat{Y}_t^c$ by $\hat{\Pi}(L)$, we have: $\hat{y}_t^c = \hat{\Pi}(L)\hat{Y}_t^c$.

   ii) Let $\hat{v}_f^c(m)$ be the smallest eigenvalue of :

$$\Phi_f^c(m) = 0.5\left[\sum_{t=2}^{T}\left(\hat{y}_t^c\hat{y}_{t-1}^{c'} + \hat{y}_{t-1}^c\hat{y}_t^{c'}\right)\right]\left(\sum_{t=2}^{T}\hat{y}_t^c\hat{y}_{t-1}^{c'}\right)^{-1} \quad \text{(A.6)}$$

   iii) Define the statistic $MQ_f^c(m) = T\left[\hat{v}_f^c(m) - 1\right]$

   C. If $H_0 : r_1 = m$ is rejected, set $m = m - 1$ and return to step A. Otherwise, $\hat{r}_1 = m$ and stop.

B. Moon and Perron panel unit root test

The simple dynamic model provided by MP consists in the following equations:

$$\begin{aligned} y_{it} &= \alpha_i + y_{it}^0 \\ y_{it}^0 &= \rho_i y_{it-1}^0 + \varepsilon_{it}, \end{aligned} \qquad \text{(B.1)}$$

where $y_{it}^0 = 0$ for all i[10].

To model the cross-correlation, BM assume that the error term ε_{it} follows a factor model:

$$\varepsilon_{it} = \beta_i^{0t} f_t^0 + e_{it}, \qquad \text{(B.2)}$$

where f_t^0 are *K*-vectors of unobservable random factor, β_i^0 are non-random factor loading coefficient vectors (also *K*-vectors), e_{it} are idiosyncratic shocks, and the number of factor *K* is possibly unknown.

Under the null hypothesis of $\rho_i = 1$ for all i=1,2,..,N, y_{it} is influenced by two components: the integrated factor $\sum_{s=1}^{T} f_t^0$ and the idiosyncratic errors $\sum_{s=1}^{T} e_s$ .With respect to the BNG test, the MP test is based only on the estimated idiosyncratic component. MP treat the factors as a nuisance parameter and propose to pool de-factored data. MP suggest removing cross-sectional dependence in the model (B.1-B.2) by multiplying the observed matrix Y of the dimension $(T \times N)$ by the projection matrix QB and compute the unbiased pooled autoregressive estimator as:

$$\rho_{pool}^{+} = \frac{tr\left(Y_{-1} Q_B Y' - NT\lambda_e^N\right)}{tr\left(Y_{-1} Q_B Y_{-1}'\right)} \qquad \text{(B.3)}$$

where *Y-1* is the matrix of the lagged observed data, $tr(\cdot)$ is the trace operator and λ_e^N is the cross-sectional average of the one-sided long run variance of the idiosyncratic errors e_{it}. The vector of factor loading $\hat{\beta}$ and the projection matrix

[10] MP also consider a model with incidental trend: $y_{it} = \alpha_{ki}' g_{kt} + y_{it}^0$, where $g_{0t} = 1$ and $g_{1t} = (1, t)'$.

Q_B are obtained by estimating the principal component of $\hat{e}'\hat{e} = \left(Y - \hat{\rho}_{pool}Y_{-1}\right)'\left(Y - \hat{\rho}_{pool}Y_{-1}\right)$, where $\hat{\rho}_{pool}$ is the OLS pooled autoregressive estimate.

Table 1. Panel unit root tests. Variable: lnG

A. Choi Panel tests			B. Pesaran panel tests			C. Bai and Ng panel tests						D. Moon and Perron panel tests		
						Idionsycratic Shocks			Common Factor $\hat{F}$					
Pm	Z	L*	P*	CIPS	CIPS*	$\hat{r}$	$Z_{\hat{e}}^{c}$	$P_{\hat{e}}^{c}$	$ADF_{\hat{F}}^{c}$	Trends $\hat{r}_1$		$\hat{r}$	t_a*	T_a^{*B}
0.130	1.510	1.136	1	-2.003	-1.929	4	-3.331	55.804		MQ_c	MQ_f	2	-0.990	-1.237
(0.448)	(0.934)	(0.087)		(0.205)	(0.257)	(BIC3)	(0.000)	(0.000)	-	4	4	(BIC3)	(0.161)	(0.108)
			2	-1.208										
				(0.995)								$\hat{r}$	t_b*	T_b^{*B}
			3	-1.228								2	-1.083	-1.048
				(0.945)								(BIC3)	(0.139)	(0.147)
			4	-0.723										
				(0.990)										

Notes: a) The PM test is the modified Fisher's inverse chi-square test (Choi, 2001). The Z test is an inverse normal test. The L* test is a modified logit test. All statistics have a standard normal distribution under H_0 when T and N tend to infinity (Choi, 2004a); b) CIPS is the mean of individual cross-sectionally augmented ADF statistics (ADF). CIPS* indicates the mean of truncated individual CADF statistics. The truncated statistics are reported only for one lag since they are always equal to not truncated one for higher lag lengths. p* denotes the nearest integer of the mean of the individual lag lengths in ADF tests; c) for each variable, the number of common factor estimated ($\hat{r}$) is estimated by the BIC_3 criterion (we use the BIC_3 criterion because N and T don't have so much difference in magnitude), with a maximum number of factor equal to 5. For idiosyncratic components $\hat{e}_{it}$, the pooled unit root statistic test are reported. $P_{\hat{e}}^{c}$ is a fisher's type statistic based on a p-valued of the individual ASF tests. Under the null hypothesis, $P_{\hat{e}}^{c}$ has a $\chi^2(2N)$ distribution whet T tends to infinity and N is fixed. $Z_{\hat{e}}^{c}$ is the standardized Choi's type test statistic. Under the null hypothesis, $Z_{\hat{e}}^{c}$ has a N(0,1) distribution. For the idiosyncratic components $\hat{F}_t$, two different cases must be distinguished: if $\hat{r}=1$, only the standard ADF t-statistic, $ADF_{\hat{F}}^{c}$ is reported. If $\hat{r}>1$ the estimated number $\hat{r}$ of independent stochastic trends in the common factors a reported (columns 4 and 5). d) t_a* and t_b* are the panel unit root test based on de-factored panel data and computed with a quadratic spectral kernel function. e) t_a^{*B} and t_b^{*B} are computed with a Barlett kernel function. f) P-values are in parenthesis.

Table 2. Panel unit root tests. Variable: lnC

A. Choi Panel tests			B. Pesaran panel tests			C. Bai and Ng panel tests						D. Moon and Perron panel tests		
						Idionsycratic Shocks			Common Factor $\hat{F}$					
Pm	Z	L*	P*	CIPS	CIPS*	$\hat{r}$	$Z_{\hat{e}}^{c}$	$P_{\hat{e}}^{c}$	$ADF_{\hat{F}}^{c}$	Trends $\hat{r}_1$		$\hat{r}$	t_a*	T_a^{*B}
										MQ_c	MQ_f			
2.459 (0.993)	2.505 (0.994)	2.293 (0.989)	1	-2.412 (0.015)	-2.412 (0.015)	2 (BIC3)	0.135 (0.447)	31.042 (0.413)	-	2	2	2	-11.291 (0.000)	-11.385 (0.000)
			2	-2.236 (0.055)								$\hat{r}$	t_b*	t_b^{*B}
			3	-1.757 (0.500)								2	-6.264 (0.000)	-6.124 (0.000)
			4	-1.576 (0.725)										

Notes: see Table 1

Table 3. Panel unit root tests. Variable: ln P

A. Choi Panel tests			B. Pesaran panel tests			C. Bai and Ng panel tests						D. Moon and Perron panel tests		
						Idionsycratic Shocks			Common Factor $\hat{F}$					
Pm	Z	L*	P*	CIPS	CIPS*	$\hat{r}$	$Z_{\hat{e}}^{c}$	$P_{\hat{e}}^{c}$	$ADF_{\hat{F}}^{c}$	Trends $\hat{r}_1$		$\hat{r}$	t_a*	T_a^{*B}
1.407 (0.079)	-0.469 (0.319)	-0.909 (0.181)	1	-2.071 (0.150)	-2.071 (0.150)	1 (BIC3)	0.296 (0.346)	32.072 (0.319)	0.905 (0.990)	-	-	1 (BIC_3)	10.098 (0.000)	-10.059 (0.000)
			2	-1.645 (0.650)								$\hat{r}$	t_b*	t_b^{*B}
			3	-1.576 (0.725)								1	-5.779 (0.000)	-5.771 (0.000)
			4	-1.639 (0.650)										

Notes: see Table 1

Table 4. Panel data Cointegration tests.

Variables	ADF	DH_p
	(Kao)	(Westerlund)
lnG, lnC, lnP	-1.859	4.285
	(0.031)	(0.000)

Notes: a) All tests assume the null hypothesis of no cointegration . b) For the ADF test, the lag order is set to one. Results are robust to different lag lengths. c) For the DH_p, a constant term is included in the regression test. For semiparametric corrections, the Bartlett kernel is employed. Bandwidth and lag orders are both set equal to the largest integer less than $4(T/100)^{2/9}$. The number of common factors are estimated using the IC_{p2} criterion with the maximum number of factors set equal to 4. The p-values for one-sided test are based on the normal distribution.

Table 5. Estimation of the model. Bai and Kao (2004) and Westerlund (2005b)

	OLS	**Adjusted OLS (two-step)**	**FM (two-step)**
lnP	-0.190	-0.190	-0.320
	[0.154]	[0.154]	[0.154]
lnC	1.012	1.037	0.676
	[0.096]	[0.102]	[0.096]

Notes: a) Bai and Kao (2004) assume that the number of factors are known. b) For the Adjusted OLS , the maximum number of factors considered for information criteria is set equal to J=5. c) The IC_{p2} information criteria is used (see Westerlund, 2005b). d) For the estimation of the long-run co-variances, we follow the recommendation of Newey and West (1994) .e) The Barlett Kernel with the bandwidth parameter set equal to the largest integer less than $4(T/100)^{2/9}$ is used. f) Standard errors are in parenthesis .

Table 6. Estimation of the model. Bai and Kao (2004) and Westerlund (2005)

	OLS	**Adjusted OLS (iterative)**	**FM (iterative)**
LnP	-0.19	-0.308	-0.324
	[0.154]	[0.168]	[0.131]
lnC	1.012	0.676	0.669
	[0.956]	[0.113]	[0.081]

Notes: see table 5.

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