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Abstract

In Lisbon the European Council proclaimed a European growth strategy. It considers an average economic "growth rate of around 3 percent as a realistic prospect for the coming years" and assigns public finances an important role in the process of achieving this goal. This paper addresses the question whether we can find empirical evidence for European countries that public finance reform affects trend growth. Focusing on time series patterns, we investigate whether there have been persistent shifts or trends in economic growth and fiscal variables over the last 40 years. In addition, we estimate a distributed lag model, which 1) indicates that government consumption and transfers negatively affect growth rates of GDP per capita over the business cycle, while public investment has a positive impact, and 2) provides robust evidence that distortionary taxation affects growth in the medium-term through its impact on the accumulation of private physical capital.

Keywords: Panel Cointegration; Public Finances; Economic Growth

JEL Classification: C22, C23, H11, O11

Non-technical summary

The European Council set forth the Lisbon Process in order to raise growth rates of output in EU countries. The Council considers "an average economic growth rate of around 3% [as] a realistic prospect for the coming years" [1]. In a follow-up to the process initiated in Lisbon, the Commission and ECOFIN Council underscored that the quality of public finances plays a crucial role for growth and employment. More specifically, they outlined the necessity of lowering the tax burden and particularly the tax wedge for low-skilled workers, making benefit systems more supportive to employment, shifting resources towards productive expenditures in health, education and physical infrastructure, and ensuring the sustainability of public finances.

The conclusions of the European Council on the European output performance leave open whether the increase in economic growth is expected to be persistent. Therefore we investigate in this paper which growth pattern could be the outcome of policy reforms in the future course of the Lisbon process. Analysing past experiences in European countries should indicate whether public finances have the potential to raise growth rates more permanently, i.e. affecting trend growth, or whether one could at best expect a transitory improvement. In analysing this aspect we built upon an approach which exploits the different time series implications of endogenous and exogenous growth theory.

As a first step we explore the time series pattern of real per-capita GDP growth rates and budgetary aggregates and subcategories of public spending and revenues. We find some persistent deterministic changes in per-capita GDP growth rates and public finances. Looking at stochastic trends, however, we also find public finance variables have generally shown persistence over time while growth rates of output appear to be fairly stable. This pattern does not exclude a long-term effect of fiscal variables per se, if expenditures and revenues have opposite long-term effects and co-move. Using recently developed panel cointegration techniques, we indeed find overwhelming evidence of cointegration between both sides of the budget, as would be expected on theoretical grounds.

We then estimate the long-run effect of fiscal policies on growth using a distributed lag model. We improve on previous studies on the nexus of fiscal policies and growth by better controlling for real business cycle effects and reverse causality, and by using effective average tax rates as proxies for taxation. The main findings are that the expenditure side of the budget appears to consistently affect long-run growth over the

[1] See Conclusions of the Presidency, page 2.

business cycle. Specifically, government consumption as well as government transfers are found to have a clear negative effect on growth, while public investment appears to positively affect growth. Less clear-cut evidence exists for a direct effect of taxation on growth. But a robust negative impact of direct taxation on physical capital accumulation is confirmed by our empirical analysis.

1) Introduction

The European Council set forth the Lisbon Process in order to raise growth rates of output in EU countries. The Council considers "an average economic growth rate of around 3% [as] a realistic prospect for the coming years" [2]. For this purpose economic policy should be geared to foster a knowledge-driven economic expansion through the spread of new technologies and higher human capital, more perfect goods and financial markets in Europe, a more employment-friendly active labour market policy and a modernisation of the welfare state as well as an investment-friendly climate brought about by regulatory changes. Many of the measures envisaged by the heads of states affect not only the regulatory setting but also public finances. In a follow-up to the process initiated in Lisbon, the Commission and ECOFIN Council underscored that the 'quality' of public finances plays a crucial role for growth and employment. More particularly they outlined the necessity to lower the tax burden and particularly the tax wedge for low-skilled workers, make benefit systems more supportive to employment, shift resources towards productive expenditures in health, education and physical infrastructure, and to ensure the sustainability of public finances.

The conclusions of the European Council on the European output performance in the future leave open which growth model actually best reflects the intentions of the heads of state. Exogenous and endogenous growth models have substantially different implications for the impact of a policy variable on economic growth. Exogenous neoclassical growth models confine the impact of fiscal policy and other policy instruments to permanently changing the level of per capita output, i.e. they alter growth rates only temporarily during the transition path to a new steady state. By contrast, endogenous growth models predict that policy variables can permanently change not only the output level but also growth rates. If the announcement of the heads of states implicitly assumes an exogenous growth framework, then we could expect output to speed up in the short- and medium run but then level off again. Conversely, the structural changes which they envisage to make Europe a more integrated, competitive and productive economy may also imply that trend growth rises permanently from currently 2-2.5 percent to 3 percent per year.

In this paper we will investigate which growth pattern would be the outcome of policy reforms in the future course of the Lisbon process. Analysing past experiences in European countries should indicate whether public finances have the potential to raise

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[2] See Conclusions of the Presidency, page 2.

growth rates more permanently, i.e. affecting trend growth, or whether one could at best expect a transitory improvement.

Numerous 'Barro-type' regression studies have attempted to test predictions of growth theories by the set of variables they incorporate. They claim to find evidence for endogenous growth if diverse policy and institutional variables included in the regressions affect long-term performance. However, these studies mainly exploit the cross sectional variation of very large samples and are not very informative if we want to focus on the European context. Using the standard set-up of these studies, relying on long-term averages of output growth, we would quickly exhaust the degrees of freedom for European countries. Moreover, this approach would not be appropriate since the European sample is rather homogeneous in several of these explanatory characteristics.

As an alternative, a small literature has emerged around this issue focusing mainly on the time series implications of the two strands of theory. If the policy variable follows a specific time series pattern, economic growth should exhibit the same behaviour under endogenous growth theory. Conversely, the time series properties of the policy-variable do not necessarily have to coincide with output growth according to exogenous growth models. Fiscal variables are a good testing ground for these hypotheses, since distortionary taxation and productive expenditures are assumed to have a permanent effect on growth rates according to endogenous growth theory, whereas they should have only level effects from a neoclassical perspective. As a consequence, endogenous growth theory would clearly render a non-zero sum of coefficients when one estimates the impact of a policy reform on output growth over the years, while exogenous growth theory would predict a sum of coefficients converging to zero. We will use these predictions as a basis to interpret the observable pattern of economic growth and public finance developments, which has not systematically been done for Europe. This is the gap in the empirical literature which our study wants to close.

The paper is organised as follows. The following section briefly describes the theoretical background and the shortcomings of the empirical evidence existing in this area. Section 3 describes the data used in the empirical exercise. We then analyse the time series properties of real per-capita output growth and public finance variables in section 4. This analysis will show that there are persistent developments in per-capita GDP growth and fiscal variables, which are broadly in line with some theoretical predictions on long-term growth. In section 5 we conduct distributed lag estimations as a more systematic check of the long-term impact of public finances. Section 6 concludes.

2) Theory and Existing Empirical Evidence

Since the mid-1980s the theoretical growth literature has above all tried to endogenize the growth rate of output in the long-run. Earlier growth models, formulated by Solow (1956) and Cass (1965) among others, conceived trend growth largely as a function of factors exogenous to public policy – such as technological progress and population growth. In their view, public policy could only affect the level of per-capita output but not have a permanent impact on the growth rate. Endogenous growth theory pioneered by the work of Romer (1986, 1990), Lucas (1988), Barro (1990) and Rebelo (1991) among others, points out mechanisms by which policy variables cannot only affect the level of output, but also steady-state growth rates. Barro (1990) constitutes one of the first attempts at endogenizing the relationship between growth and fiscal policies. He distinguishes four categories of public finances: productive vs. non-productive expenditures and distortionary vs. non-distortionary taxation. Government spending is considered productive if it enters the private production function by contributing directly to output. Otherwise, it is considered unproductive and does not exert any lasting effect on the growth rate. Taxation is distortionary if it affects the investment decision, and hence output growth. This is, above all, the case for direct income and profit taxation. Otherwise taxes, such as consumption taxes, are considered non-distortionary, except for the case when households face the endogenous choice of labour or leisure.

We present a simple sketch of the Barro-model in order to show that both productive public expenditure and distortionary taxation can affect long-run output growth[3]. We assume that the population of consumers is normalised to one. Consumers both consume and produce final output according to the following production function:

$$y = Ak^{1-\gamma}g^{\gamma} \tag{1}$$

where k stands for privately accumulated physical capital and g is productive government expenditure that directly enters the production process. It is assumed that the government budget constraint is balanced in every period and is given by:

$$g + G = \tau \cdot y + T \tag{2}$$

where G represents other government expenditure that does not directly enter the production function as input, T represents lump-sum taxation and τ is a proportional tax on output that distorts the investment decision. Consumers maximise their intertemporal utility function that is given by $\int_0^{\infty} e^{-\rho t}\frac{c^{1-\sigma}-1}{1-\sigma}dt$ subject to the standard budget constraint.

[3] For more details on the model, see Barro (1990) and Barro and Sala-I-Martin (1995).

ρ represents the time preference rate at which future consumption is discounted and σ is the elasticity of intertemporal substitution. The growth rate of consumption and output in steady state takes the form:

$$\frac{\dot{c}}{c} = \frac{\dot{y}}{y} = \frac{1}{\sigma}\left[(1-\tau)(1-\alpha)A^{\frac{1}{1-\alpha}}\left(\frac{g}{y}\right)^{\frac{\alpha}{1-\alpha}} - \rho\right] \tag{3}$$

Equation 3 shows that productive government expenditure as a share of output positively affects long-run growth while distortionary taxation has a negative impact on growth. Neither unproductive expenditure nor lump-sum taxation affect output growth in steady state. From this model, we see that fiscal variables from both sides of the budget constraint matter for growth, and the failure to include both productive government expenditures and distortionary taxation in regressions would lead to mis-specified models.

Jones (1995) constitutes the first attempt at exploiting time series properties to test exogenous vs. endogenous growth theory. He starts with the simple argument that according to endogenous growth theory, permanent shifts in certain policy variables should have a permanent effect on the growth rate of the economy. Hence, if growth rates in the US and other OECD countries exhibit no large persistent changes, the underlying policy variables should also either not show large persistent changes or the persistent movements in these variables must be off-setting. Using the conventional ADF test, he finds considerable evidence for a stochastic trend in the data-generating process for total investment and producer durables investment in a large share of countries[4]. He also looks at the development of R&D and total factor productivity, finding again persistent changes in R&D expenditure. However, no such persistent changes are found for total factor productivity growth. He then estimates a distributed lag model in order to assess the actual impact of a permanent shift in investment and R&D expenditures. From this exercise he concludes that the data apparently refute predictions of endogenous growth theory and the macroeconomic-policy variables under consideration do not appear to exert any permanent effect on growth. Instead, the effect is rather short-lived, dying out after a few periods.

[4] As pointed out by Jones (1995), any macroeconomic variable expressed as a share of GDP such as the private investment share or total revenue share of GDP, cannot be driven by a pure unit root process, since such variables appear to be bounded between zero and one, and a stochastic process characterised by a pure unit root would cross such a boundary sooner or later. Then he continues arguing that the investment share can be easily driven by a stochastic trend within the interval from zero to one. Thus it is necessary to bear such a point in mind when testing below for unit roots in variables expressed as shares of GDP.

Following the same line of reasoning as Jones (1995), Romero de Ávila (2002) is unable to refute the empirical validity of AK models for a wide sample of countries over a period of more than forty years. By making use of panel methods to analyse the time series properties of the series, the author finds that both the investment share of GDP and per capita output growth rates are stationary. A significant long-run impact from productive physical investment on growth is becomes apparent when estimating distributed lag models. The results are found to be robust with respect to the definition of physical investment, the use of instrumental variables, as well as the omission of outliers. These results stand in stark contrast to those by Jones, in all likelihood due to the low power of individual ADF tests to reject a false null of a unit root.

Karras (1999) largely follows the approach of Jones (1995), focusing on the effect of taxation on per capita GDP growth. He applies the approach to a panel of 11 OECD countries, finding that the real GDP growth rate is generally stationary, while the null of a unit root cannot be rejected for the total and direct tax rates in most of the countries in his sample. He concludes that adjustments of tax rates cannot be associated with permanent changes of real GDP growth, unless permanent changes in taxes are cancelled out by permanent changes in other policy variables. But he does not investigate this final possibility by including the expenditure side of the budget in his analysis. The same holds true for Evans (1997) who employs the distributed lag approach using a sample of 92 countries when he analyses the impact of government consumption on growth, not obtaining clear-cut evidence supporting the endogenous growth paradigm.

Kocherlakota and Yi (1997) test whether taxes or public investment have any permanent effect on output growth, based on time series for up to 100 years for the US and 160 years for the UK. Thus they incorporate both sides of the budget into their analysis and find that predictions of exogenous growth theory are usually rejected when taxes and public investment are included in the econometric model. However, they do not formally test for the co-movement of policy variables. This also holds true for Kneller et. al (1999) and Bleaney et al. (2001) when estimating the long-run effect of public finances on growth for the OECD countries. They use data for functional categories of central government revenues and expenditures in order to compute more precise aggregates of productive and unproductive expenditures as well as distortionary and non-distortionary taxation. As theory would predict, they find a significant impact from productive expenditures and distortionary taxation on growth while no discernible impact is found for non-distortionary taxation and unproductive expenditures.

3) Data

As the previous short review of the empirical studies in this area of research has shown, the existing evidence clearly supports endogenous growth predictions of a long-term impact when both sides of the budget are taken into account, but evidence is still incomplete. Except for the study of Bleaney et al. (2001) the sample of countries and the budgetary categories used have been very selective. Bleaney et al. (2001) incorporate most European economies in their sample and include all budget items, but they focus on central government data only. The drawback of this approach is, of course, that ideally one would look at general government figures. First, overall government activity and not only central government activity should count from an economic point of view. Second, general government provides a more homogeneous data set than central government, which may vary strongly according to the organisation of national and subnational authorities.

Therefore, we use data for general government outlays and revenues in all EU member states from 1960 to 2001 (Commission AMECO data set, Autumn 2002). All time series are computed in logs and fiscal variables are measured as shares of GDP. Budgetary aggregates are classified according to an economic criterion rather than functionally. This is fairly unproblematic with respect to taxation, because the classification of direct taxation on property and income, on the one hand, and indirect taxation on imports and production on the other, largely reflects the theoretical distortionary/non-distortionary classification. For public expenditures the link is less immediate. Evidently, public capital formation could be counted as productive expenditures. However, even this is somewhat questionable since the definition of gross fixed capital formation includes investments directly supporting private production, such as necessary roads, but also inefficient 'pork barrel' projects. For government consumption similar arguments hold. It comprises wage payments going to teachers and professors, i.e. they are investments in human capital, as well as salaries and purchases for the social security system, which Bleaney et al. (2001) assume to be unproductive. In addition, it also comprises expenditures on health and education, which are clearly of productive use. In other words, empirical evidence on the impact of these spending categories jointly evaluates the validity of the theoretical predictions and assesses the productive vs. non-productive content of the expenditure flow under consideration. Data to compute real GDP per capita as well as the private investment share of GDP are obtained from the OECD Economic Outlook.

4) The Time Series Properties of Growth and Public Finances

To assess the potential impact of fiscal policies on growth we first take a look at the time series pattern of our dependent and independent variables. The conjecture that permanent shifts in policy variables should be associated with a permanent shift in the growth pattern, if endogenous growth theory holds, is compatible with different time series patterns. Therefore, we first search for deterministic long-run movements and then for possible persistent stochastic processes.

4.1) Deterministic Trends and Breaks

Figure 1 presents the growth rates for real GDP per capita in EU member states from 1961 to 2001. From the charts there is not a clear time series pattern evident which would hold for all countries. Growth rates in several countries, such as France, Spain, Germany, Greece, Italy and Sweden, have slowed down over the last few decades. In Ireland and Luxembourg medium-term growth has however picked up over this time period. Moreover, it is not fully clear whether the slowdown of per-capita GDP growth is a smooth trend or associated with a break. Most countries show a fall in the growth rate around the early 1970s, associated with the first oil price shock. For Greece, Portugal and Spain, most obviously, this break could have initiated a period of sustained lower growth. Individual countries – in particular Germany and Finland – also experienced severe disruptions in the early 1990s, which, at least in the German case, seem to be associated with lower trend growth over the subsequent decade.

More formal evidence for these patterns is presented in Table 4. In the first column we show different estimates for deterministic trends existing in EU countries. The first row shows a trend estimate imposing a common mean and trend coefficient for all countries. The second row shows the within estimate of the trend coefficient (i.e. allowing the intercept to vary) and the third row a mean group estimate of the trend coefficient that is allowed to vary across countries. In the fourth row we present the results of a test for a common break point. A short explanation of the estimated equations and the mean group estimator can be found in Appendix 1. The first fact to note is that the estimated coefficient for a deterministic trend in per-capita GDP growth carries a negative sign and is highly statistically significant. Moreover, the size of the coefficient and the significance level is the same for all three estimates, whether we impose common intercepts and trends or not.

Turning to public finances, Figures 2 and 3 show the development of total spending and revenues in EU Member States over the last four decades. All countries are marked by a clear increase in public spending up to the early 1980s. This trend flattens thereafter or is even reversed. In Belgium, Ireland, Luxembourg and the Netherlands, the reversal sets in in the 1980s, while it is of a more recent nature in most other countries. Overall public revenues show a similar but often less pronounced increase in European countries in the 1960s and 1970s. This trend then also flattens in most cases, but is not reversed. Only the

United Kingdom shows a trend of slightly decreasing overall revenues from the early 1980s to mid-1990s.

Aggregate spending and revenues are obviously inaccurate measures of productive expenditures and distortionary taxes, which should have long-run effects on growth. Table 4 therefore presents the same tests as mentioned above for budgetary aggregates as well as different economic spending and revenue categories, which might have an impact on growth according to the theoretical framework spelled out above. The second column confirms the long run upward trend in government spending. This was mainly driven by transfers and to a somewhat lesser extent by government consumption. Interestingly, public investment shows the opposite development, as indicated by the negative and statistically significant coefficient for the whole panel. Looking at whether one can also identify a common break point in the trend of public spending, the last row indicates that spending growth (measured in first differences) indeed decelerated, and the common break point identified by our method is 1983. This is directly related to the breaks that can be found for the increasing transfer payments and consumption expenditures in the early 1980s. Looking at public investment however, the decline in public investment according to these estimates accelerated after 1972. Both break points are statistically significant, although the overall explanatory power of the model for public investment is relatively low. Public revenues and distortionary taxation are marked by upward trends, being somewhat larger for distortionary taxation. The breakpoint, after which the trend flattened, is dated to 1984 for total revenues and to 1977 for distortionary taxation.

The overall trend developments are largely compatible with predictions from growth theory. Both policy variables and per-capita GDP growth show long-run developments. The long-term decline of public investment and rising distortionary taxation are both compatible with the lower trend growth apparent in our estimates. The downward shift in the GDP per-capita growth rate during the first oil price shock is still in line with the accelerated reduction of public investment after 1972. Similarly, data would suggest that the accelerated increase of distortionary taxation up to 1977 contributed to a downward adjustment of growth in the mid-1970s, although the upward trend of taxation flattened thereafter. Breakpoints for other revenue and spending categories, for which the productive or distortive quality is less clear, are somewhat harder to reconcile with the per-capita GDP growth pattern.

4.2) Stochastic Trends

Given the observable patterns in Figures 1 to 3, our policy variables in particular may not only exhibit deterministic but also persistent stochastic trends. To test this hypothesis we carry out panel unit root tests since unit root tests for individual time series suffer from a lack of power for the number of observations available in our data set. We employ the Im, Pesaran and Shin (IPS) and the Breitung Test and use different specifications, with and without heterogeneous trends. An explanation of why these tests and specifications are chosen as well as a brief description of the test statistic is provided in Appendix 2.

Table 5 presents the results of the unit root tests for real per-capita growth, current revenues and its main subcategories (direct taxation, social security contributions and indirect taxation) as well as total expenditures, transfers, government consumption and public investment.[5] For real per-capita output the null hypothesis of a unit root cannot be rejected with any test while it can be safely rejected for the real GDP growth rate. In short, for the sample of European countries we can assume that per-capita GDP follows an I(1) pattern, while real GDP per-capita growth is mean-reverting.

The results for fiscal variables are also clear. As far as the revenue side of the budget is concerned, the IPS-test indicates a unit root in all specifications for total revenues. The same conclusions emerge from the Breitung test. For distortionary taxation specifications with deterministic trends also indicate a persistent stochastic trend. This result seems to be mainly driven by the direct tax component, since the results for social security contributions are rather mixed. While the IPS test rejects the null of a unit root at the 1% level when trends are included and at 10% in the specifications without trends, the Breitung test does not reject it.

With regard to the expenditure side of the budget, total expenditure appears to be driven by a stochastic trend, since we could only reject the null at the 10% for the specification augmented with four lags and without trends. The same evidence arises for all the expenditure subcategories including investment. Government consumption, government transfers and public investment all contain a unit root. At the bottom of Table 5 we also present the unit root statistics for the variables in first-differences. The test results clearly indicate that no fiscal variable is integrated of second order.

Overall the picture emerging from these tests is that for the entire panel real per-capita output growth is stationary. For aggregate public finances and most subcategories the evidence is also quite clear with the exception of social security contributions, where the IPS and the Breitung tests point to opposite conclusions. Given that public investment is generally considered as productive expenditure and direct taxation and social security contributions as distortionary taxes, this result could be considered as a challenge to endogenous growth theory. Since persistent changes in fiscal categories do not appear to be accompanied by persistent changes in GDP per capita growth rates, there is no support

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[5] We include all economic categories for the sake of completeness, although only some results will be discussed in this section. But the time series properties are also relevant for the following sections.

for endogenous growth predictions according to the logic expressed by Jones (1995) and Karras (1999).

4.3) Cointegration of Expenditures and Revenues

The diverging time series pattern of per-capita GDP growth and our fiscal policy variables could still be reconciled if another policy variable with an offsetting persistent effect on growth existed. In other words, two conditions have to be met: first, the variable has to co-move with the policy instrument under consideration and, second, it should exhibit a persistent growth effect according to endogenous growth theory. The obvious candidate here is to look at the opposite side of the budget, since any expenditure increase has to be financed, and this may lead to a higher excess burden of taxation.

It can be shown that the intertemporal budget constraint does indeed imply a cointegration relationship between revenues and expenditures (see Afonso 2003, Santos Bravo and Silvestre 2002, Trehan and Walsh 1988, Bohn 1998). However, whether this relationship exists for our sample and whether it could explain the observable pattern of economic growth is still an empirical question. First, growth theory focuses on productive expenditures and distortionary taxation. Thus it is empirically unclear whether the cointegration relationship holds for relevant spending and revenue items. For example, higher public investment could be financed through non-distortionary consumption taxes. Second, the intertemporal budget constraint has to hold over an infinite horizon and it is binding under the assumption that the economy operates efficiently. The implications over a finite horizon are not clear ex ante, particularly if the economy operates inefficiently and growth rates exceed interest rates. Under these circumstances, countries can engage in a “deficit gamble” even for an extended period of time (Ball et al. 1998, O’Connell and Zeldes 1988).

Therefore, we conduct panel cointegration tests for different combinations of spending and revenue aggregates or sub-categories. In order to identify long-run relationships, we present the results of the seven panel cointegration tests proposed by Pedroni for the specifications with and without heterogeneous deterministic trends[6]. We opt for normalising on the variable standing for the revenue side of the government budget constraint, without implying, of course, that the direction of causality is running from the spending side of the budget constraint to the revenue side.[7]

[6] The panel unit root and cointegration analysis was carried out with the NPT1.3 package developed by Chiang, and Kao (2003) in addition to some routines kindly provided by Peter Pedroni.

[7] Empirical evidence on Granger-causality for expenditures and revenues indicates different patterns running unidirectional from expenditures to revenues or vice versa, or being bi-directional in some countries (Belessiotis, 1995)

Results in Table 6 corroborate the existence of a long-run relationship between revenues and spending. First, total spending as a share of GDP appears to clearly cointegrate with total current revenues when deterministic trends are included in the cointegrating vector, as one would expect with a budget constraint. Similar cointegrating patterns are found for the long-run relationship between total expenditures and revenue subcategories in the specification with deterministic trends.

Given this result we would expect 'big ticket items' to co-move with the other side of the budget. Indeed, we find clear indications that total transfers cointegrate with all revenue subcategories, in particular with distortionary taxation.[8] Regarding total government consumption as a share of GDP, we find relatively fewer indications of cointegration with the other side of the budget, and in general only the pooled and mean group ADF t-statistics can reject the null of no cointegration. If anything, the development of indirect taxation is relatively closely aligned with the persistent stochastic trend in government consumption. Finally, there is firm evidence that public investment cointegrates with total revenue as well as with distortionary taxation in the specification without trends.[9] This is in stark contrast to the cointegrating results for other expenditure figures, where the inclusion of trends is important to capture the co-movement between revenue and expenditure figures.

Table 7 presents the estimates of the long-run coefficient for the variables entering the cointegrating vector. The long-run estimates should provide evidence for the co-movement between fiscal categories from both sides of the budget. We base our inferences mainly on the FMOLS estimators, which correct the standard OLS for the bias induced by the endogeneity and serial correlation of the regressors. The group-mean FMOLS moreover is preferable to the pooled FMOLS since it allows for more flexible hypotheses testing. Other estimates are nevertheless presented in the table mainly as a robustness check for the reader. The cointegrating vector is again normalised on the revenue category.

Looking at these relationships from the expenditure side, Table 7 shows that the long-run coefficient of the relationship of aggregate revenues with total expenditures is around 0.7 for all estimators and is robust against the inclusion of time effects. This implies that an increase in total government spending is under-compensated by an increase in total

8 Somewhat less evidence of a long-run cointegrating relationship between government transfers and aggregate revenues is found since the null of no cointegration could only be rejected with three tests at the 5% or better (the pooled ADF, the mean group Philips-Perron and the mean group ADF t-statistics).

9 The unit root test indicated an I(1) process irrespective of the inclusion of a trend. Thus it is not clear that the cointegration test including a trend is the more relevant statistic, even if we find a deterministic trend in the previous section.

revenues.[10] We tested for a one-to-one relationship between total revenues and total expenditures, rejecting the existence of a proportional relation at the 1% significance level. This finding can be partly explained by the deficit bias leading to a continuous debt build-up during the 1970s and 1980s in many European countries.[11]

Furthermore, the coefficient of the relationship between transfers and total revenues falls from 0.46 to 0.37 when time dummies are included. The coefficients for distortionary taxation, direct taxation and social security contributions show similar decreases in magnitude when we include time dummies, indicating that part of the co-movement between these budget categories is accounted for by factors which are invariant across EU countries, such as common supply or demand shocks. The coefficients of government consumption with respect to total revenues and distortionary taxation are around 0.8 without time effects and 0.2 when accounting for common shocks. The cointegration relationship with indirect taxation that was apparent in the above tests renders a coefficient of 0.4 and is significant when no time dummies are included. The long-run elasticity of public investment to total revenues and distortionary taxation appears to be 0.15 for the specification with time dummies. Cointegration coefficients however become negative and very small if the time effects are not included.

In short, this exercise has shown a strong cointegration relationship between public revenues and expenditures. This holds particularly true for budget aggregates, but also applies to sub-categories. Moreover, the cointegrating vectors generally have the expected sign, showing that expenditures and revenues co-move in the same direction. To the extent that their persistent component may have opposite effects on growth, they would therefore cancel each other out. An interesting case is government consumption since there is strong evidence for cointegration with indirect taxation. If government consumption comprises a persistent growth-enhancing component, this would not be off-set fully by distortionary tax developments.

5) The Impact of Public Finances on Long-term Growth – A Distributed Lag Test

The previous exercises searching for deterministic processes have shown that per-capita GDP growth rates and fiscal policy variables are marked by persistent, long-run

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[10] Again, when we talk about undercompensating movements in revenues as a result of expenditure changes, we do not mean causality going from expenditures to revenues.

[11] This is in line with the findings in Afonso (2003) and Santos Bravo and Silvestre (2002) who conduct cointegration tests for individual EU member states. Interestingly however, the coefficient of direct taxation on total expenditures is greater than one when time dummies are included, pointing to an overcompensating movement of direct taxation as a response to changes in aggregate expenditures.

developments. Real per-capita GDP shows a declining trend in the growth rate, or possibly a downward shift of growth after the early 1970s. Conversely, public finance variables follow some persistent upward or downward trends and shifts which could be compatible with this pattern. When we look at stochastic processes, fiscal variables also reveal a persistent component, while GDP growth is mean-reverting. This pattern is still compatible with the assumption that the persistent impact of productive expenditures and distortionary taxation cancel each other out. But even under this condition, fiscal policies may nevertheless have a short- or medium term impact and influence growth over the business cycle. In this section we will therefore analyse more systematically whether fiscal variables affect economic performance over a cycle.

5.1) Estimation Procedure

The long-term effect of fiscal policy on growth can be estimated using a distributed lag approach, controlling for both sides of the government budget. Thus a simple model of the following form will be estimated:

$$\Delta y_{it} = \sum_{j=0}^{I} \beta_{gj} g_{it-j} + \sum_{j=0}^{I} \beta_{\tau j} \tau_{it-j} + \theta_t + \varepsilon_{it} \tag{8}$$

Here y indicates the log of per capita output, and g and τ represent government expenditures and revenues respectively. This equation can theoretically be used to test exogenous vs. endogenous growth theories. As Evans (1997) and Kocherlakota and Yi (1997) show, exogenous growth theory implies

$$\sum_{j=0}^{I} \beta_{gj} g_{it-j} = \sum_{j=0}^{I} \beta_{\tau j} \tau_{it-j} = 0$$

as the lag order goes to infinity Conversely, endogenous growth theory implies for productive expenditures

$$\sum_{j=0}^{I} \beta_{gj} g_{it-j} > 0$$

and for distortionary taxation

$$\sum_{j=0}^{I} \beta_{\tau j} \tau_{t-j} < 0 \, .$$

In other words, the sum of coefficients has to be different from zero for a sufficiently large lag order if endogenous growth predictions are to be valid. It is not clear *ex ante* what constitutes the right lag order in this context. Even transitory changes to the new equilibrium state after a fiscal reform can expand over several years. Thus, we cannot sensibly discriminate between the two theories, but only confirm whether public finances have a consistent impact on growth over the cycle and affect trend growth. Therefore we use a lag-length equal to 8 since spectral analysis tends to indicate a business cycle of 6 to 8 years for European countries (see Boutevillain et al. 2001). Moreover, this specification is in line with the literature in the public finance field (Bleany et al 2001).

Equation 8 can be rewritten as a function of the lag operator as follows:

$$\Delta y_{it} = A(L)g_{it} + B(L)\tau_{it} + \theta_t + \varepsilon_{it} \tag{9}$$

where *A(L)* and *B(L)* represent two lag polynomials with unit roots outside the unit circle. We re-parameterise equation 9 in line with Jones (1995) in order to separate long-run from short-run effects as follows:

$$\Delta y_{it} = A(1)g_{it} + \mathrm{C(L)}\,\Delta g_{it} + B(1)\tau_{it} + D(L)\Delta\tau_{it} + \theta_t + \varepsilon_{it} \tag{10}$$

where *C(L)* and *D(L)* are (p-1)th-order lag-polynomials such that:

$$c_{is} = -\sum_{j=s+1}^{p} a_{ij}$$

$$d_{is} = -\sum_{j=s+1}^{p} b_{ij}$$

where s = 1, ………, p-1.

In sum, the coefficients for *A(1)* and *B(1)* capture the long-run effect of government spending and revenue categories on growth, while the first-difference terms will capture short-run interactions between fiscal policies and growth.

A source of concern of these distributed lag growth models is the likely endogeneity of fiscal policies as a result of governmental responsiveness to current and future growth prospects. This problem implies that current as well as lagged fiscal policies may be correlated with the error term in equation 9. This could also be the result of other common factors, which as the previous exercise made clear are important in the European context. This problem is tackled by including time dummies and, following the procedure proposed by Li (2002), leads along with the distributed lags for the policy variables in our models[12].

[12] This practice is similar to the approach by Stock and Watson (1993) who compute the DOLS estimator in order to correct for the endogeneity of regressors and serially-correlated errors by using leads of the first-differenced regressors.

To clarify the approach, let us consider the case when the error structure in equation 9 that may be correlated with the regressors takes the following form:

$$\varepsilon_{it} = G(L)g_{it+q} + H(L)\tau_{it+q} + v_{it} \tag{11}$$

where *G(L)* and *H(L)* are polynomials of order 2q. We assume that the new error term v_{it} is uncorrelated with leads and lags of productive expenditures and distortionary taxation, as otherwise the reverse causality problem through business cycle effects would remain. It is also assumed that for large enough values of q, the correlation between ε_{it} and the expenditure and taxation terms are zero beyond q leads and lags. It is further assumed that cyclical shocks can only affect fiscal adjustments and growth in the short-run, since we do not expect the existence of a long-run relation between ε_{it} and fiscal policies. As such, equation 11 can then be rewritten as $\varepsilon_{it} = G(1)g_{it} + G'(L)\Delta g_{it+q} + H(1)\tau_{it} + H'(L)\Delta\tau_{it+q} + v_{it}$, and since *G(1)* and *H(1)* are assumed to be zero, we have that: $\varepsilon_{it} = G'(L)\Delta g_{it+q} + H'(L)\Delta\tau_{it+q} + v_{it}$. By substituting ε_{it} into the expression for growth, it renders:

$$\Delta y_{it} = A(1)g_{it} + I(L)\Delta g_{it+z} + B(1)\tau_{it} + K(L)\Delta\tau_{it+z} + \theta_t + v_{it} \tag{12}$$

where *I(L)* and *K(L)* are lag polynomials of order 2z equal to *G'(L)* and *H'(L)* respectively when z >0 to account for the number of leads in the polynomials, and *I(L)* and *K(L)* are *C(L)+G'(L)* and *D(L) + H'(L)* respectively for z≤0, as given by the current and lagged terms in the polynomials. For our computational purposes we include eight lags and five leads[13]. Finally, we also include the private investment rate in the regressions since capital accumulation is theoretically the prime engine of growth.

5.2) Estimation Results

5.2.1) Public Finances and Growth

Table 8 reports the results for aggregate revenues and expenditures and different sub-components.[15] In models (1) to (3) we estimate the overall effect that government size has on growth. Model (1) controls for both sides of the budget constraint to render a

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[13] The results appear fairly robust to different lag and lead-lengths. In order to keep a reasonable number of useable observations we set the number of leads to five. Panel studies often average the data over five-year periods to cancel out cyclical fluctuations.

[14] Later in the section, we will analyse the link between private physical investment and distortionary taxation.

[15] We have estimated the same models without leads. Throughout the analysis, the results are generally robust against the inclusion of leads, which reaffirms the fact that what we are capturing are long-run growth effects. The results are available from the authors upon request.

significant coefficient on total expenditures equal to –0.045 while revenues are insignificant. Obviously, the high collinearity between both aggregates with a correlation coefficient larger than 0.9 may inflate the standard errors and lead to parameter instability for the revenue coefficient in model 1. Therefore we estimate the same model with each budget aggregate separately. [16] Model (2) shows that the coefficient for current revenues becomes negative once we drop total expenditures from the regression. In model (3) we only include total expenditures rendering a coefficient equal to –0.031, which captures the benefits of spending minus the cost of taxation in addition to the reduction of growth due to deficit financing. According to model (3), a 1% increase in the total expenditure share ceteris paribus will lead to a cumulative decrease in growth by 3.1% over the long-run. Thus any positive growth effect of productive expenditures is more than cancelled out by its negative financing implications, even if it is at least partially deficit financed. Overall, our estimates indicate that the coefficient value of the costs of distortionary taxation minus the benefits from productive expenditures is around 0.02 and 0.04.

In order to disentangle whether the balance of benefits and costs looks different for individual spending categories, we disaggregate total expenditures that comprise all productive and unproductive government outlays into government consumption, transfers and public investment. In models (4) to (10), we analyse the sign and size of the effect stemming from these major economic categories of government expenditures after controlling for different revenue categories. For the rest of the analysis we also control for private investment since we are interested in the effect that individual fiscal policies have on growth beyond their effect on the investment decision.

From regression (4), we can see that the spending categories are highly significant when we include total revenues as the financing control. The net effect of government consumption and total transfers is clearly negative, yielding coefficients of –0.035 and –0.038 respectively, while public investment has a positive coefficient equal to 0.016. Therefore, a 1% increase in the public investment share of GDP would bring about a rise in growth of 1.6% over the long-run. Private investment is also significant at the 5% level, implying that a 1% increase in the private investment share of GDP would cause a 1.1% increase in growth. These estimates for expenditure categories remain fairly similar and mostly statistically significant when we control for different revenue components, although the estimates are not always very stable.

[16] We also report the t-statistics and the significance level of the estimates, although these do not allow proper inference since the GDP growth rate is stationary and policy variables are I(1). However, it should be noted that the results are largely in line with those reported later on, where we control properly for both sides of the budget. To the extent that spending and revenue items are cointegrated, the t-statistics can be used for statistical inference.

According to the simple theoretical model presented at the beginning, major spending items could be non-productive and therefore at worst neutral with respect to per-capita growth. Finding again the negative and very robust relationship already apparent for total expenditures, although we have now properly accounted for the financing side plus any possible fall in public investment that could result from different factors. It could reflect either a financing cost that is not captured by our aggregate tax measure, the costs of deficit financing in terms of growth, or a growth-reducing impact which is not apparent in our simple theoretical model, where labour supply is not a crucial variable. Social benefits or government wages may however reduce labour supply in the private sector and thereby undermine growth.[17]

The results for taxes are less robust by comparison.[18] Total revenues expressed as a share of GDP appear marginally significant with a counterintuitive positive coefficient. Models (5) to (8) control for an individual revenue subcategory. Direct taxation is found to be significant and with a positive coefficient. The same holds for indirect taxation although the impact is less significant. Social security contributions are significant and negatively correlated with growth. In models (9) and (10) we simultaneously control for several revenue subcategories in the same regression to make sure our results are not driven by omitted variable bias. In these specifications the negative impact of social security contributions is not statistically significant any more.

These results indeed appear hard to reconcile with the underlying growth model. The counterintuitive positive effect from direct taxation may result from a measurement error considering that we use a very rough proxy, i.e. the revenue to GDP ratio, to capture the distortionary effect of taxation on growth (see also Slemrod 1995). Economic theory suggests that marginal tax rates affect the investment and labour-leisure decisions of workers, and in turn long-run growth. Considering the difficulties associated with a consistent computation of marginal tax rates to capture these marginal effects for the EU-15 (as the system of exemptions and scales used to tax different types of income vary widely across countries), we have used revenue shares to capture the average tax burden. Nevertheless, changes in the tax system leading to an increase in public revenues could, for example, result from a lowering of tax rates through a simultaneous broadening of the

[17] The negative coefficient which we find is much in line with the results reported in de la Fuente (1997) for industrialized countries. De la Fuente also discusses the productivity and investment channels through which government expenditures could have a negative externality for growth. These are however partly captured by our tax controls. In addition, he provides evidence for strongly decreasing returns to public investment. We have estimated a non-linear specification of our model to test for this conjecture, but could not find evidence in this direction.

[18] This is in line with other empirical studies in this area (see de la Fuente 1997, Myles 2000).

tax base. This is the direction that several reforms of the systems of personal and corporate income taxation have taken in recent years[19].

The use of effective tax rates has been suggested in the literature to at least approximately capture the average negative effect that taxation may have on growth. These effective tax measures may thus be a better proxy than the revenue shares of GDP. Therefore, we run a few more growth regressions using effective tax rates on labour, capital and consumption goods as a check of our results. We obtained the data from Martínez-Monguay (2001) who computed these measures for the EU-15 for the period 1970-2001. The results are presented in Table 9. As can be observed in models (11) to (14), neither the effective tax rates on labour nor on capital are found to be significantly related to growth. Only the effective tax rate on consumption is found to be significant and positive for growth, probably reflecting the positive benefits from a shift from direct to indirect taxation.

5.2.2) Taxation and Private Investment

These results for public revenues that at first sight appear disappointing may indicate the need to focus more on theoretical elements when analysing the impact on growth. In fact, endogenous growth theory suggests that distortionary taxation affects the investment decision, and provided private investment exerts a positive impact on growth, distortionary taxation in turn affects growth. In order to check for this transmission channel of physical capital accumulation, we run distributed lag regressions with private investment as the dependent variable on different revenue categories, also controlling for aggregate government outlays.

In Table 10 we present the results when using revenue to GDP ratio as proxies for tax effects, while in Table 11 we present some more distributed lag regressions for private investment using the effective tax rates. These results appear to be more encouraging. Model (15) makes use of total revenue as a proxy for taxation, also controlling for total expenditure. Considering the high collinearity between the two, it is not surprising that neither of them is significant. Model (16) drops total expenditure and total revenue becomes highly significant and the coefficient carries the expected sign. The long-run coefficient implies that an increase in overall taxation by 10% would bring about a fall in private investment by around 3%. In models (17) to (22) we control for different revenue subcategories. As in model 15, when the total expenditure share of GDP is included along with direct taxation (model 17), social security contributions (model 21) and distortionary taxation (model 19), all long-run elasticities become insignificant. This stems from the

[19] See European Commission (1999, 2000, 2001, 2002) for details.

high correlation between aggregate expenditures and the revenue subcategories[20]. Once we drop aggregate expenditures, models (18) and (22) render significant coefficients on direct taxation of around -0.14. Model (20) yields a coefficient on distortionary taxation equal to -0.116[21].

In Table 11, we further investigate the robustness of the link between investment and taxation by using the effective tax rates from Martínez-Monguay (2001). Models 23 to 28 give a clear message. Capital taxation consistently and negatively affects investment. The coefficient takes on a negative value of around 0.4, implying that a 1% increase in capital taxation would cause a fall in the private investment share by 0.4%. We could not however find any significant effect from the effective tax of labour income on private investment.

Our results on the link between effective tax rates, long-run growth and investment appear in line with Mendoza et al. (1997). They could not find any statistically significant negative impact from taxation on growth when using data averaged over five-year periods to control for the cycle, while some evidence could be found when using annual data. As a result, they argue that the negative effect that factor taxation may have on growth may be of a short-run nature. But their main result is based on the highly significant negative impact that effective tax rates on capital exert on private investment rates.

Overall, our results point to the existence of a significant long-run effect from aggregate government expenditure and its main subcategories on growth over the business cycle. The size of the public sector as well as government consumption and transfers negatively affects long-run growth. This may reflect the fact that the average size of the public sector in the EU is above its optimal level. Conversely, public investment has a positive impact on growth which indicates the likely gains in economic performance from shifting welfare expenditure to productive investment. Furthermore, we find that taxation affects growth through its impact on private capital accumulation.

6) Conclusions

The Lisbon Process assigns a prominent role to public finance reform in order to foster economic growth. The main purpose of our analysis is therefore to shed some light on the relation between public finances and growth in the EU15. Most importantly, this requires

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[20] The correlation coefficients of total expenditures with total revenues, direct taxation and distortionary taxation equal 0.923, 0.758 and 0.887 respectively.

[21] In the regressions without leads, the same coefficient equals -0.176, which points to the fact that business cycle effects may play a role.

determining whether public finances provide policy instruments contributing to higher trend growth, or whether they can at best be expected to have a short-run impact on economic performance. Following the approach of some studies in this field that exploit the time series properties of the data, we find some persistent deterministic changes in per-capita GDP growth rates and public finances. Looking at stochastic trends, however, we also find public finance variables have generally shown persistence over time while growth rates of output appear to be fairly stable. This pattern does not exclude a long-term effect of fiscal variables per se, if expenditures and revenues have opposite long-term effects and co-move. Using recently developed panel cointegration techniques, we did indeed find overwhelming evidence of cointegration between both sides of the budget, as would be expected on theoretical grounds.

We then estimate the long-run effect of fiscal policies on growth using a distributed lag model. We improve on previous studies on the nexus of fiscal policies and growth by better controlling for real business cycle effects and reverse causality as well as by using better proxies for taxation. The main findings are that the expenditure side of the budget appears to consistently affect long-run growth over the business cycle. Specifically, government consumption as well as government transfers are found to have a clear negative effect on growth, while public investment appears to positively affect growth. Less clear-cut evidence exists for a direct effect of taxation on growth. But a robust negative impact of direct taxation on physical capital accumulation is confirmed by our data. These results stand in contrast to previous studies on this issue which did not appropriately take into account the financing relations implied by the intertemporal budget constraint.

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Table 1: Notation

LYPC	Gross Domestic Product in per capita terms nominated in constant dollars
DLYPC	Growth rates of GDP per-capita
LPRINV	Private physical investment share of GDP
LTREV	Total current revenues as a share of GDP
LTEXP	Total expenditures as a share of GDP
LPI	Total public investment as a share of GDP
LTR	Total transfers as a share of GDP
LG	Government consumption spending as a share of GDP
LTDIR	Total direct taxation as a share of GDP
LSSC	Social security contributions as a share of GDP
LTDIST	Total distortionary taxation as a share of GDP
LTIND	Total indirect taxation as a share of GDP
LTL	Effective labour tax rate
LTK	Effective capital tax rate
LTC	Effective consumption tax rate

Note: All variables are expressed in log-levels.

Table 2: Descriptive Statistics

	OBS.	MEAN	STD. DEV.	MIN	MAX
LYPC	630	16145.601	7507.463	3619.8172	66398.89
LINV	600	20.415272	4.644229	5.0138457	35.411656
LTREV	597	40.628469	9.71107	17.438311	62.859245
LTDIR	544	12.925482	5.763888	2.3692704	30.644902
LSSC	544	11.398384	4.924931	1.4631681	21.056033
LTDIST	544	24.323866	7.227372	7.6919585	39.626596
LTIND	544	13.601302	3.274826	5.8751693	30.644902
LTEXP	597	42.81849	10.2771	17.385859	70.075973
LG	536	17.323588	4.313626	7.8456798	28.858792
LTR	544	15.284265	5.057451	3.0295877	28.473822
LPI	544	3.2873537	1.052914	1.0336889	6.4685041
DLYPC	615	2.835303	2.683282	-11.0494	11.55296

Notes: All variables are expressed as percentages of GDP, except growth rates of GDP p.c.

Table 3: Correlation Matrix

	LY	LINV	LTREV	LSSC	LTIND	LTDIR	LTDIST	LTEXP	LG	LTR	LPI	DLY
LYPC	1.000											
LINV	-0.498	1.000										
LTREV	0.765	-0.377	1.000									
LSSC	0.330	-0.121	0.308	1.000								
LTIND	0.376	-0.167	0.567	-0.251	1.000							
LTDIR	0.672	-0.371	0.864	-0.062	0.575	1.000						
LTDIST	0.790	-0.348	0.953	0.462	0.378	0.818	1.000					
LTEXP	0.670	-0.358	0.923	0.342	0.490	0.758	0.887	1.000				
LG	0.462	-0.162	0.764	-0.017	0.594	0.719	0.658	0.745	1.000			
LTR	0.670	-0.373	0.820	0.542	0.316	0.581	0.862	0.876	0.439	1.000		
LPI	-0.095	0.046	-0.044	0.058	-0.051	-0.031	-0.062	-0.062	-0.151	-0.074	1.000	
DLYPC	-0.175	0.137	-0.306	-0.175	-0.054	-0.220	-0.332	-0.411	-0.330	-0.363	0.054	1.000

Note: The matrix shows the correlations of relevant fiscal variables (measured as share of GDP).

Figure 1.

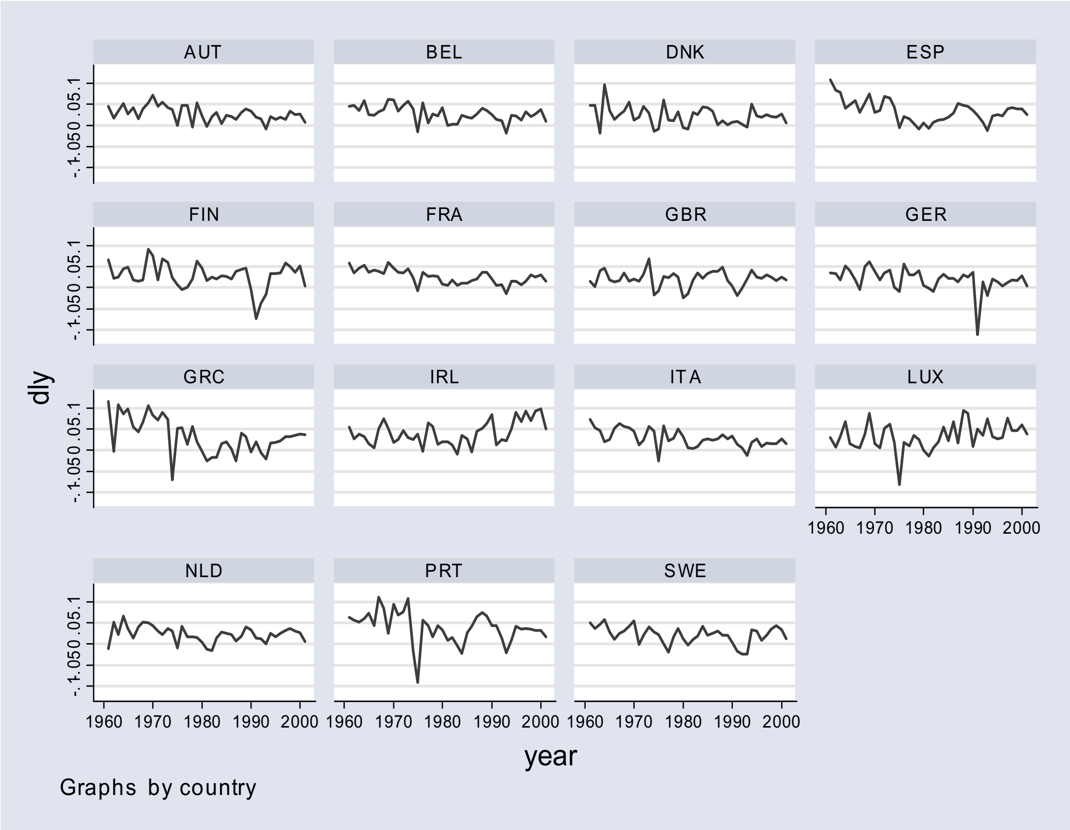

Figure 2.

Figure 3.


AUT
BEL
DNK
ESP
FIN
FRA
GBR
GER
GRC
IRL
ITA
LUX
NLD
PRT
SWE
ltrev
3
3.5
4
4.5
1960
1970
1980
1990
2000
year
Graphs by country

Table 4: Deterministic Time Series Patterns

	DLYPC	LTEXP	LTR	LG	LPI	LTREV	LTDIST
deterministic trend (common coefficient and intercept)	-0.0005*** (0.00013)	0.015*** (0.0015)	0.018*** (0.003)	0.010*** (0.002)	-0.014*** (0.003)	0.014*** (0.001)	0.017*** (0.003)
deterministic trend (fixed effects, common coefficient)	-0.0005*** (0.0001)	0.015*** (0.001)	0.015*** (0.002)	0.010*** (0.0009)	-0.013*** (0.002)	0.013*** (0.0008)	0.015*** (0.001)
deterministic trend (mean group estimate)	-0.0005*** (0.0002)	0.015*** (0.0019)	0.015*** (0.003)	0.011*** (0.001)	-0.014*** (0.005)	0.014*** (0.002)	0.015*** (0.002)
endogenous break	-0.023*** (0.002) (1974)	-0.029*** (0.003) (1983)	-0.042*** (0.006) (1983)	-0.018*** (0.005) (1982)	-0.039*** (0.015) (1972)	-0.015*** (0.003) (1984)	-0.024*** (0.004) (1977)
F-Statistic	*117.07*	*72.57*	*56.39*	*14.41*	*6.48*	*31.37*	*34.68*

Note: For explanations see Appendix 1. Figures in parenthesis are standard errors. Asterisks indicate statistical significance at the one percent level.

Table 5: Panel Unit Root Tests

Variables	Z-bar IPS Test 2 Lags		Z-bar IPS Test 4 Lags		Breitung
	No trend	Trend	No Trend	Trend	
LYPC	4.30	3.14	3.87	2.32	2.328
LCREV	1.97	-0.73	0.06	-1.01	0.400
LTDIR	1.75	2.04	-0.43	2.47	-0.617
LSSC	-1.49*	-2.60***	-1.60*	-2.77***	0.686
LTDIST	0.48	-0.31	-1.94**	-0.16	1.167
LTIND	2.23	0.43	2.67	1.44	-1.571*
LTEXP	1.26	1.38	-1.51*	-1.00	3.033
LPI	1.45	2.61	0.80	2.85	-0.540
LTR	0.68	-0.06	1.16	-0.12	4.379
LG	1.02	0.38	1.60	2.24	-1.373
DLYPC	-5.28***	-5.54***	-2.27***	-2.19***	-11.273***
DCREV	-9.50***	-8.26***	-5.64***	-3.96***	-12.791***
DLTDIR	-8.93***	-8.41***	-3.63***	-3.50***	-13.713***
DLSSC	-8.77***	-6.94***	-5.90***	-3.06***	-13.125 ***
DLTDIST	-9.70***	-8.69***	-5.96***	-5.02***	-12.932***
DLTIND	-8.24***	-6.91***	-5.52***	-5.15***	-12.685***
DLTEXP	-6.96***	-5.40***	-3.18***	-1.96*	-10.625***
DLPI	-6.80***	-5.28***	-2.12***	0.15	-16.602***
DLTR	-6.96***	-5.46***	-4.57***	-2.84***	-10.970***
DLG	-8.33***	-6.72***	-4.67***	-3.64***	-13.110***

Note: Time dummies were included in all ADF specifications. *, ** and *** imply rejection of the null of unit root at the 10%, 5% and 1% level of significance respectively.

Table 6: Panel Cointegration Results

	Country Dummies	Country Dummies and Trend	Country Dummies	Country Dummies and Trend	Country Dummies	Country Dummies and Trend	Country Dummies	Country Dummies and Trend
Dep. Variable	**Indep. Variable**		**Indep. Variable**		**Indep. Variable**		**Indep. Variable**	
LTREV	**LTEXP**		**LPI**		**LTR**		**LG**	
panel v-stat	1.447	4.241***	-2.142	-0.360	-0.802	1.043	-0.074	-0.226
panel rho-stat	-0.445	-2.676***	-0.412	1.001	0.622	-0.765	-0.779	0.894
panel pp-stat	-0.189	-4.085***	-1.535**	-0.140	0.475	-2.360	-1.356*	-0.094
panel adf-stat	-0.398	-5.750***	-1.493**	-0.049	0.328	-3.359***	-1.898**	-0.315
group rho-stat	0.075	-1.470*	-0.186	1.450	1.388	0.397	0.813	1.151
group pp-stat	-0.006	-4.031***	-2.199**	-0.004	0.633	-1.879**	-0.375	-0.050
group adf-stat	-1.269	-7.121***	-2.479***	-1.903**	-0.363	-6.489***	-1.661**	-1.509*
LTDIR	**LTEXP**		**LPI**		**LTR**		**LG**	
panel v-stat	1.752**	3.262***	-0.794	1.305*	0.676	2.285**	0.813	0.873
panel rho-stat	-1.181	-2.094**	-0.924	-0.440	-0.234	-1.452*	-1.014	-0.137
panel pp-stat	-1.143	-3.092***	-1.921**	-1.457*	-0.069	-2.680***	-1.415*	-1.063
panel adf-stat	-2.239**	-6.573***	-2.261**	-1.409*	-0.289	-5.104***	-2.871***	-2.119**
group rho-stat	-0.585	-0.943	-0.355	0.065	0.708	-0.787	-0.707	0.252
group pp-stat	-0.895	-2.877***	-2.170**	-1.392*	0.446	-2.704***	-1.532*	-0.937
group adf-stat	-3.954***	-11.657***	-3.633***	-3.372***	-0.688	-10.979***	-3.966***	-4.519***
LSSC	**LTEXP**		**LPI**		**LTR**		**LG**	
panel v-stat	-0.129	1.249	-0.753	-1.517	-0.298	2.141**	-0.866	0.473
panel rho-stat	0.414	-0.645	-1.182	1.722	0.205	-0.515	0.799	0.847
panel pp-stat	-0.229	-1.809**	-2.550***	0.628	-0.413	-1.800**	-0.301	-0.371
panel adf-stat	-0.309	-3.235***	-2.745***	0.530	-0.947	-3.859***	-0.398	-0.967
group rho-stat	0.445	0.403	0.450	2.688	0.578	0.263	1.687	1.540
group pp-stat	-0.350	-1.420*	-2.008**	1.159	-0.335	-1.726**	0.063	-0.229
group adf-stat	-4.673***	-7.101***	-2.885***	0.157	-3.34***	-5.713***	-1.049	-4.903***
LTDIST	**LTEXP**		**LPI**		**LTR**		**LG**	
panel v-stat	1.422*	3.765***	-1.560	0.684	-0.070	3.098***	-0.572	-0.457
panel rho-stat	-0.889	-2.282**	-0.864	0.668	0.455	-1.450*	-0.173	1.515
panel pp-stat	-1.099	-3.477***	-2.365***	-0.677	0.406	-2.661***	-1.090	0.267
panel adf-stat	-1.985**	-8.215***	-2.603***	-1.300*	0.417	-4.752***	-1.769**	-1.298*
group rho-stat	-1.068	-1.377*	0.322	1.567	0.815	-0.536	0.705	1.739
group pp-stat	-1.469*	-3.463***	-2.255**	-0.084	0.078	-2.608***	-0.866	0.413
group adf-stat	-4.994***	-11.341***	-2.638***	-1.750**	-1.181	-7.637***	-2.304**	-5.280***
LTIND	**LTEXP**		**LPI**		**LTR**		**LG**	
panel v-stat	0.849	1.870**	0.645	1.213	-0.522	2.250**	0.576	0.454
panel rho-stat	0.861	-1.420*	-1.068	-1.239*	0.579	-0.807	-0.155	-0.599
panel pp-stat	1.267	-2.614***	-1.425*	-2.787***	0.373	-2.265**	-0.246	-1.939**
panel adf-stat	1.605	-5.815***	-2.381***	-5.786***	0.177	-6.381***	-0.388	-2.659***
group rho-stat	1.816	-0.289	-0.416	-0.067	1.631	-0.298	1.519	0.609
group pp-stat	1.987	-2.364***	-1.172	-2.355***	1.121	-2.834***	0.958	-1.319*
group adf-stat	1.351	-10.883***	-7.803***	-12.426***	-0.273	-18.580***	-0.231	-6.006***

Note: The cointegrating vector is normalised on the revenue category. *, ** and *** imply rejection of the null of unit root at the 10%, 5% and 1% level of significance respectively.

Table 7: Panel Cointegrating Vectors

Dep. Variable	Indep. Variable		Indep. Variable		Indep. Variable		Indep. Variable	
	Country Dummies	Country Dummies and Time Dummies	Country Dummies	Country Dummies and Time Dummies	Country Dummies	Country Dummies and Time Dummies	Country Dummies	Country Dummies and Time Dummies
LTREV	**LTEXP**		**LTR**		**LG**		**LPI**	
Mean Group OLS	0.67***	0.63***	0.46***	0.32*	0.71***	0.26*	-0.04	0.12
	(5.65)	(4.57)	(2.81)	(1.9)	(3.57)	(1.84)	(-1.47)	(1.35)
1-to-1 Rel Test	(5.65) ***	(4.57) ***						
MG FMOLS	0.66***	0.71***	0.46***	0.37***	0.75***	0.26***	-0.01***	0.15***
	(31.25)	(26.6)	(15.88)	(11.26)	(20.6)	(10.48)	(-8.45)	(-57.42)
1-to-1 Rel Test	(-13.52) ***	(-13.84) ***						
Pooled Unweighted FMOLS	0.72***	0.71***	0.46***	0.37**	0.72***	0.39	-0.04	0.2***
	(5.8)	(5.1)	(3.69)	(2.53)	(3.78)	(1.59)	(-0.37)	(-9.28)
1-to-1 Rel Test	(-2.29)**	(-2.1)**						
LTDIR								
Mean Group OLS	0.75***	1.06***	0.6*	0.43	0.84**	0.35	0.04	0.25
	(3.21)	(2.62)	(1.95)	(1.07)	(2.39)	(1.28)	(-0.66)	(1.18)
MG FMOLS	0.74***	1.27***	0.61***	0.46***	0.92***	0.35***	0.1***	0.32***
	(17.82)	(15.89)	(10.99)	(6.3)	(14.2)	(7.42)	(-3.5)	(7.96)
Pooled Unweighted FMOLS	0.98***	1.08***	0.59***	0.52	1.00***	0.65	0.03	0.39**
	(4.3)	(3.04)	(2.94)	(1.61)	(3.31)	(1.31)	(0.12)	(2.39)

Table 7: Panel Cointegrating Vectors (continued)

Dep. Variable	**Indep. Variable**		**Indep. Variable**		**Indep. Variable**		**Indep. Variable**	
	Country Dummies	Country Dummies and Time Dummies	Country Dummies	Country Dummies and Time Dummies	Country Dummies	Country Dummies and Time Dummies	Country Dummies	Country Dummies and Time Dummies
LSSC	**LTEXP**		**LTR**		**LG**		**LPI**	
LSSC								
Mean Group OLS	0.87***	0.53**	0.64***	0.34**	0.83***	0.05	-0.14	0
	(4.62)	(2.35)	(4.26)	(1.98)	(2.59)	(0.29)	(-1.31)	(0.32)
MG FMOLS	0.87***	0.52***	0.65***	0.39***	0.81***	0.01	-0.13***	0.02***
	(26.23)	(13.06)	(24.35)	(11.73)	(14.95)	(1.22)	(-7.62)	(2.6)
Pooled Unweighted FMOLS	0.81***	0.38*	0.61***	0.38**	0.69***	0.05	-0.19	0.05
	(3.72)	(1.71)	(3.77)	(2.19)	(2.29)	(0.18)	(-1.38)	(0.44)
LTDIST								
Mean Group OLS	0.76***	0.65***	0.58***	0.34***	0.76***	0.22	-0.05	0.11
	(5.8)	(3.03)	(3.61)	(2.07)	(2.95)	(0.83)	(-1.18)	(1.04)
MG FMOLS	0.76***	0.74***	0.6***	0.38***	0.8***	0.2***	-0.03***	0.15***
	(32.64)	(18.06)	(20.54)	(12.27)	(17.1)	(4.58)	(-6.61)	(7.08)
Pooled Unweighted FMOLS	0.83***	0.66***	0.56***	0.41**	0.78***	0.32	-0.08	0.2*
	(4.7)	(3.64)	(4.02)	(2.44)	(3.16)	(1.06)	(-0.54)	(1.96)

Table 7: Panel Cointegrating Vectors (continued)

Dep. Variable	Indep. Variable		Indep. Variable		Indep. Variable		Indep. Variable	
	Country Dummies	Country Dummies and Time Dummies	Country Dummies	Country Dummies and Time Dummies	Country Dummies	Country Dummies and Time Dummies	Country Dummies	Country Dummies and Time Dummies
LTIND	**LTEXP**		**LTR**		**LG**		**LPI**	
LTIND								
Mean Group OLS	0.22	-0.11	0.16	0.56	0.37	-0.09	-0.09	0.04
	(0.56)	(0.75)	(0.4)	(0.53)	(0.72)	(0.23)	(-0.08)	(0.95)
MG FMOLS	0.16**	-0.15***	0.12**	0.7***	0.41***	-0.1	-0.12	0.02***
	(2.43)	(4.12)	(1.85)	(3.6)	(4.41)	(1.17)	(-1.1)	(5.18)
Pooled Unweighted FMOLS	0.28	0.2	0.16	0.06	0.4	0.24	-0.08	0.06
	(0.99)	(0.5)	(0.7)	(-0.06)	(1.29)	(0.69)	(-0.55)	(0.4)

Note: The cointegrating vector is normalised on the revenue category. The row labelled by 1-to-1 Rel Test relates to testing the existence of long-run proportionality between total government revenues and aggregate expenditure. T-statistics are given in parenthesis below the estimates. *, ** and *** imply rejection of a zero long-run elasticity at the 10%, 5% and 1% level of significance respectively.

Table 8: Public Finances and Long-term Growth

	1	2	3	4	5	6	7	8	9	10
LPRIVINV				0.0110**	0.0137***	0.0069	0.0079	0.0045	0.0135**	0.0157***
				(2.023)	(2.969)	(1.393)	(1.551)	(0.956)	(2.552)	(2.976)
LTREV	0.0112	-0.0175**		0.0437*						
	(0.717)	(-2.217)		(1.819)						
LTEXP	-0.0450**		-0.0312***							
	(-2.318)		(-4.065)							
LTDIR					0.0147***				0.0129**	0.0190***
					(3.291)				(2.378)	(3.429)
LSSC						-0.0078***			-0.0028	0.0031
						(-3.549)			(-1.008)	(0.849)
LTDIST								0.0191		
								(1.520)		
LTIND							0.0169*			0.0185*
							(1.985)			(1.782)
LG				-0.0352**	-0.0257***	-0.0148**	-0.0180**	-0.0127	-0.0256***	-0.0353***
				(-2.499)	(-3.225)	(-2.261)	(-2.087)	(-1.604)	(-3.026)	(-3.557)
LTR				-0.038***	-0.026***	-0.008	-0.016***	-0.033***	-0.022***	-0.029
				(-3.058)	(-5.574)	(-1.560)	(-3.082)	(-3.168)	(-3.544)	(-4.560)
LPI				0.016***	0.0174***	0.0229***	0.0166***	0.0178***	0.0192***	0.0072
				(2.681)	(3.608)	(3.948)	(2.720)	(3.162)	(3.760)	(1.145)
R2	0.609	0.500	0.575	0.741	0.758	0.737	0.734	0.742	0.770	0.790

Note: The dependent variable is given by the growth rate of GDP per capita. Estimations are carried out with pooled data for the EU-15. The sample period ranges from 1960-2001. All regressions include unreported time effects. All t-statistics were computed using Newey-West standard errors to correct for heteroskedasticity and serially correlated errors characteristic of distributed lag models. The table reports the long-term coefficients A(1) and B(1) of equation 12 in the text. The T-statistics are given in parenthesis. *, ** and *** imply the rejection of the null at the conventional significance levels of 10%, 5% and 1% respectively. The F-statistic tests for the significance of all variables included in the regression. See table 1 for the notation.

Table 9: Effective Tax Rates and Growth

	11	12	13	14
LPRIVINV	0.0075	0.0094	0.0109*	0.0041
	(1.351)	(1.556)	(1.829)	(0.803)
LTEXP	-0.0477***	-0.0333*	-0.0281***	-0.0155
	(-3.693)	(-1.860)	(-2.923)	(-0.664)
LTl	0.0066	0.0036		-0.0061
	(0.894)	(0.336)		(-0.439)
LTk	0.0052	0.0089	0.0072	
	(0.801)	(1.171)	(0.834)	
LTc	0.0199***			
	(3.410)			
R2	0.713	0.678	0.651	0.591

Note: See Table 8 with the only difference that the sample period goes from 1970-2001 since data on effective tax rates were not available before 1970. The table reports the long-term coefficients A(1) and B(1) of equation 12 in the text. The T-statistics are given in parenthesis. *, ** and *** imply the rejection of the null at the conventional significance levels of 10%, 5% and 1% respectively.

Table 10: Public Finances and Private Investment

	15	16	17	18	19	20	21	22
LTREV	-0.776	-0.315**						
	(-1.336)	(-2.225)						
LTEXP	0.697		0.192		0.233		0.471	
	(1.060)		(0.793)		(0.529)		(1.144)	
LTDIR			-0.147	-0.143***			-0.270	-0.142***
			(-1.510)	(-3.444)			(-1.596)	(-3.331)
LSSC							-0.035	0.029
							(-0.642)	(0.926)
LTDIST					-0.209	-0.117*		
					(-0.793)	(-1.861)		
R2	0.282	0.201	0.402	0.328	0.394	0.332	0.458	0.395

Note: The dependent variable is the private investment share of GDP. For the rest see table 8' footnote. The table reports the long-term coefficients A(1) and B(1) of equation 12 in the text. The T-statistics are given in parenthesis. *, ** and *** imply the rejection of the null at the conventional significance levels of 10%, 5% and 1% respectively.

Table 11: Effective Tax Rates and Private Investment - Specification with 5 Leads

	23	24	25	26	27	28
LTEXP	0.3420		0.4073		0.1888	
	(1.038)		(1.018)		(0.984)	
LTL	-0.0998	-0.0123	-0.1311	0.0350		
	(-0.651)	(-0.144)	(-0.683)	(0.378)		
LTK	-0.4050**	-0.3652***	-0.3915**	-0.3567***	-0.4229**	-0.4268***
	(-2.575)	(-3.407)	(-2.288)	(-2.970)	(-2.472)	(3.371)
LTC	0.0645	0.1401				
	(0.708)	(1.579)				
R2	0.512	0.996	0.468	0.415	0.442	0.382

Note: See table 10. The sample period goes from 1970-2001 since data on effective tax rates were not available before 1970. The table reports the long-term coefficients A(1) and B(1) of equation 12 in the text. The T-statistics are given in parenthesis. *, ** and *** imply the rejection of the null at the conventional significance levels of 10%, 5% and 1% respectively.

Appendix 1: Deterministic Time Series Pattern

The deterministic trend pattern is first estimated using the following equation:

$$y_{it} = \alpha + \delta \mathrm{T}_i + e_{it} \tag{A.1}$$

y is the variable under consideration for country *i* at time *t* and *T* presents the time trend. In the table we present the estimates of δ and New-West Standard errors allowing for serial correlation.

Under the second specification α is indexed to country *i*. In other words, we estimate a fixed effects model where the intercept is allowed to vary across countries. δ then represents the common within estimator. In the third specification the above equation is estimated freely for each country and then the mean group estimator δ_{MG} is computed. Pesaran et al. (1996) have shown that the mean group estimator

$$\delta_{MG} = \frac{1}{n}\sum_{i=1}^{n}\delta_i$$

is asymptotically consistent, although less efficient than a pooled estimate if the homogeneity assumption holds. A consistent (non-parametric) estimate of the variance of the mean group coefficient vector, which allows for the computation of the test statistic, is provided by

$$\hat{V}(\delta_{MG}) = \frac{1}{n(n-1)}\sum_{i=1}^{n}(\delta_i - \delta_{MG})(\delta_i - \delta_{MG})'$$

For the shift mean test we follow the approach employed by Jones (1995) using the following regression model:

$$\Delta y_{it} = \alpha_i + \delta I_{[t>t^*]} + e_{it} \tag{A.2}$$

where *I* is a dummy variable that takes the value of one for t > t*. The reported estimate is the δ coefficient for the year maximizing the F-statistic for the test δ = 0.

Appendix 2: Panel Unit Root Tests

We utilise the panel unit root test deveoped by Im, Pesaran and Shin (2003). They present a group-mean ADF-t statistic for testing the null of non-stationarity versus stationarity for at least a - fraction of panel members. Compared to time series unit-root tests for individual countries, the pooling of information dramatically increases the number of observations, and hence the power of the test. Our panel specification will be of the form:

$$\Delta y_{it} = \alpha_i + \delta_i t + \theta_t + \gamma_i y_{it-1} + \sum_{i=2}^{pi}\beta_{ij}\Delta y_{i,t-i+1} + \varepsilon_{it} \tag{A.3}$$

where p_i is the required degree of lag augmentation to make the residuals white noise, α_i and $\delta_i t$ represent the specific cross-section fixed effects and deterministic trends respectively, and θ_t denotes the time dummies used to account for cross-correlations and interdependencies across different members of the panel. This test also allows for heterogeneous autoregressive coefficients on $y_{i,t-1}$ under the alternative hypothesis.

In applying the IPS test we have to decide whether to use a homogeneous or heterogeneous degree of augmentation in the individual ADF regressions for each member of the panel. According to Maddala and Wu (1999), one is implicitly restricted to use homogeneous lag-truncation for each individual ADF test, as only in this case the tables presented by IPS with small-sample adjustment factors can be used. As a result, we will compute this test using a homogeneous lag-truncation of two and four, since in most cases the longest lag-truncation found in individual ADF-t tests was in general four or lower than four[22]. The results in general appear to be robust to different lag-lengths used to augment the ADF specifications.

The main strengths of the test, compared with others such as the Levin and Lin test, is that γ_i is allowed to differ across countries and only a fraction of members of the panel is required to be stationary under the alternative hypothesis. However, there are obvious drawbacks. First, the cross-section units must be independently distributed for the validity of the IPS test. This may be a strong assumption taking into account the likely correlation across countries. Second, the IPS test suffers from an enormous decrease in power when country-specific trends are included in the specification as a result of the bias correction applied to the t-statistics (Baltagi and Kao, 2000; Breitung, 2000)[23]. To address the issue of cross-correlation, we include time dummies that will account for all common shocks affecting all members of the panel in a given period and in turn for the cross-correlation that may arise as a result.

The second problem is addressed by using a panel unit root test that is not sensitive to the inclusion of deterministic trends. Breitung (2000) has proposed a panel unit root test which employs unbiased t-statistics. This is achieved by transforming the variables which no longer require any small-sample bias correction. By allowing for heterogeneous

[22] The degree of augmentation for the individual time series were computed following the general-to-specific step-down procedure by which it is necessary to remove insignificant lag-differenced terms until the last term is significant at conventional levels of significance. Individual ADF-results are available from the authors upon request.

[23] In fact, as shown by IPS (1997), the correct specification of the order of augmentation of individual ADF-t tests becomes even more important in the presence of deterministic trends in the regression. They indeed show that the t-bar group-mean ADF test is less sensitive to over-specification of the order of augmentation of the individual ADF regressions than to under-specification. They also demonstrate that the t-bar test appears to be more favourably affected in terms of power by a rise in T than in N.

deterministic trends and short-run dynamics across countries without the need of bias adjustment, the Breitung test is characterised by a considerable gain in power compared to the IPS test. Moreover, the Breitung test does not seem to be sensitive to the lag-length of the augmenting terms in first-differences as opposed to the other tests.

Appendix 3: Cointegration of Fiscal Variables

The first attempts at analysing cointegration within a panel setting applied panel unit-roots tests directly to estimated residuals. That analysis is in line with the two-step procedure proposed by Engle-Granger (1987) for individual time series. Nevertheless, Pedroni (1995) shows the inappropriateness of this procedure. In contrast to the time series case, the lack of exogeneity of the regressors will give rise to off-diagonal elements in the asymptotic covariance matrix that are idiosyncratic across countries. As a result, these elements will not disappear even asymptotically, thereby rendering inconsistent estimates of the cointegrating vectors.

Therefore, we will make use of seven panel cointegration tests presented by Pedroni (1997,1999), since he determines the appropriateness of the tests to be applied to estimated residuals from a cointegrating regression after normalising the panel statistics with correction terms. These tests allow for variations in the degree of permissible heterogeneity across countries and in the extreme case, by pooling only the multivariate unit roots information, it will be possible to leave the potential cointegrating vectors entirely heterogeneous across countries. Although our sample of countries is fairly homogeneous, we thus do not base our testing procedure on this restriction[24].

We now briefly outline the approach developed by Pedroni. The procedures proposed by Pedroni makes use of estimated residuals from the hypothesised long run regression of the following form (Pedroni, 1999):

$$y_{it} = \alpha_i + \delta_i t + \gamma_t + \beta_{1i} x_{1it} + \beta_{2i} x_{2it} + ... + \beta_{Mi} x_{Mit} + e_{it} \qquad (A.4)$$

where M is the number of regressors, and N and T are the number of cross-section units and time observations respectively. This can be seen as a fixed effects model, where α_i is the country specific intercept, and γ_t represents a set of time dummies common to all members of the panel. There might be some cases when it is also appropriate to include

[24] As Pedroni (1996) argues, by incorrectly imposing homogeneity of the cointegrating vectors, this could imply that the null of no cointegration might not be rejected when there is indeed cointegration.

deterministic time trends, $\delta_i t$, specific to each individual panel member. The coefficients, β_{Mi}, are allowed to differ across individuals[25].

In line with traditional time series analysis, determining whether or not the relationships under consideration cointegrate is equivalent to demonstrating whether the estimated errors, $\hat{e}_{it}$, in equation A.6 are stationary. This can be done by establishing whether ρ_i equals one in the following equation:

$$\hat{e}_{it} = \rho_i \hat{e}_{i,t-1} + u_{it} \tag{A.5}$$

According to Pedroni (1997,1999), four of the seven proposed statistics are based on pooling along the within-dimension and three of them are pooled along the between-dimension. For the within-dimension statistics the test for the null of no cointegration is implemented as a residual-based test of the null hypothesis of H_0: $\rho_i=1$ for all *i*, versus H_1: $\rho_i=\rho<1$ for all *i*. By contrast, for the between-dimension statistics the null of no cointegration is H_0: $\rho_i=1$ for all *i* versus H_1: $\rho_i<1$ for all *i*, thus allowing for an additional source of heterogeneity across countries under the alternative hypothesis.

Following Pedroni's terminology, we will refer to the within-dimension based statistics as pooled cointegration statistics and the between-dimension as mean-group cointegration statistics[26]. The first of the pooled statistics is similar to the non-parametric variance ratio statistic developed by Phillips and Ouliaris (1990) for the time-series case. The second is a panel version of the Phillips and Perron rho-statistic. The third is also non-parametric and is analogous to the Phillips and Perron t-statistic. The fourth is a parametric t-statistic similar to the ADF t-statistic for time series analysis[27].

With respect to the group-mean cointegration statistics, we first have a group-mean rho-statistic, and the remaining two tests are similar to the PP and ADF t-statistics, respectively[28]. Pedroni rescales each of the seven test statistics so that they are distributed as a standard normal distribution. The standardisation of the cointegration statistics can be expressed as:

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[25] This constitutes the main difference of the tests presented by Pedroni (1999) with respect to those by Pedroni (1995) where the long-run coefficients were assumed to be homogeneous across countries.

[26] The term mean-group widely used in the panel literature relates to the computation of statistics (and estimators) by averaging across individual statistics (and estimators) for all member of the panel.

[27] Phillips and Ouliaris (1990) demonstrated that the PP t-statistic and the ADF t-statistic are asymptotically equivalent. Pedroni (1997) showed that the same holds for the pooled and the group-mean panel cointegration PP and ADF t-statistics.

[28] A complete explanation of which circumstances the different tests perform better under can be found in Pedroni (1997).

$$\frac{\kappa_{NT} - \mu\sqrt{N}}{\sqrt{\nu}} \Rightarrow N(0,1)$$

where κ_{NT} is the standardised form of the test statistic with respect to N and T[29]. The values of the mean (μ) and the variance (ν) are tabulated in Pedroni (1999). The values of the normalised statistics are to be compared to the critical values implied by a one-tailed standard normal distribution. For the panel variance ratio statistic, large positive values imply that the null of no cointegration is rejected. For the other six statistics, large negative values imply rejection of the null.

Pedroni (1996, 2000) proposes a group-mean and a pooled fully modified ordinary least squares estimator (henceforth FMOLS) to estimate the long-run cointegrating vector once the variables under consideration have been found to be non-stationary and cointegrated. These estimators allow for endogenous regressors without implying endogeneity bias and control for the idiosyncratic serial correlation patterns present in the residuals of individual members of the panel. Finally, the advantage of the mean-group FMOLS t-statistic over the pooled FMOLS t-statistic is that it enables us to test the null hypothesis H_0: $\beta_i = \beta_0$ for all i cross-section units versus the alternative hypothesis H_1: $\beta_i \neq \beta_0$ for all cross-section units. As a result, the values for β_i are not constrained to be homogeneous across members of the panel under the alternative. In contrast, the pooled FMOLS t-statistic tests the null H_0: $\beta_i = \beta_0$ for all i against the alternative H_1: $\beta_i = \beta_1 \neq \beta_0$, so that all cross-sections units are constrained to have the same coefficient β_1 under the alternative.

29 This scaling factor is $T^2 N^{3/2}$ for the pooled variance ratio statistic, $TN^{1/2}$ for the pooled rho (ρ), $TN^{-1/2}$ for the mean-group rho, and $N^{-1/2}$ for both the mean-group PP and ADFt-statistics.

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