Austrian Government Expenditures: "Wagner's Law" Or "Baumol's Disease"?

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ABSTRACT

Government expenditures have grown in Austria during most of the 20th century. In this paper, we present empirical evidence for this growth process and analyze some of its possible reasons. In particular, two prominent theoretical explanations for public sector growth are tested for Austria: first, "Wagner's Law" hypothesizing a positive income elasticity of demand for public goods, and second, "Baumol's Cost Disease", relating public sector growth to above-average cost increases in the public sector as compared to the private sector. The empirical evidence confirms the importance of the "Cost Disease" for Austria but cannot confirm the validity of "Wagner's Law". Business cycles influence government expenditures in the short run, while a number of variables suggested by public choice theories – except for fiscal illusion – do not significantly influence the growth of the public sector in Austria.

1. Introduction

The role and the importance of the public sector in industrialized economies have changed during the recent decades. For a long time, its activities were mostly considered as positive, such as, for instance, the importance of public infrastructure for economic growth, of transfers for reducing an uneven distribution of income and wealth, and of fiscal policies for income stabilization and employment. This was especially true for Austria with its long tradition of enlightened absolutism, viewing government as benevolent and well-informed. However, the predominant assessment of the public sector in Austria has recently become more critical. As in other European countries and elsewhere (cf. Tanzi and Schuknecht, 2000), it is increasingly asked whether limitations to the growth of the public sector are to be established in order to enhance the competitiveness of the Austrian economy.

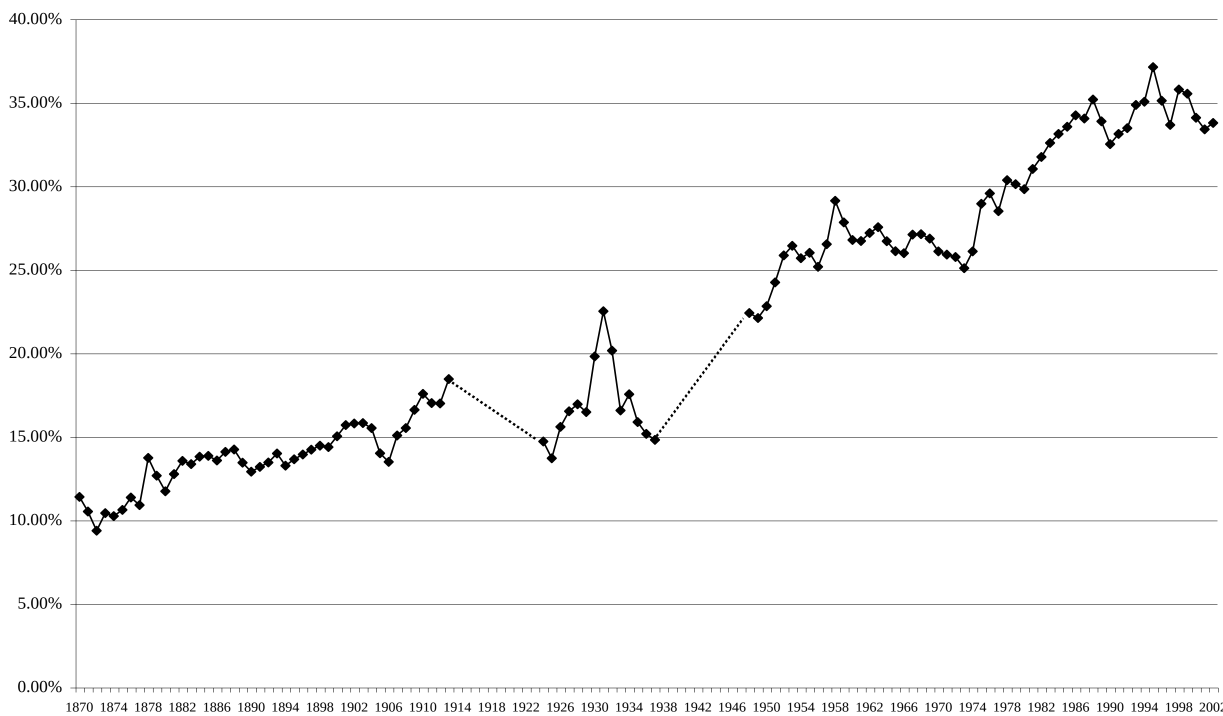
In view of this political debate in Austria, an exploration of the development and the determinants of government expenditures is of particular importance. In this paper, we describe briefly the historical development of government expenditures since 1870 in Section 2. In Section 3, we present "Wagner's Law" and other explanations of the growth of the public sector such as "Baumol's Cost Disease", the "Displacement Effect", and arguments from public choice theories. In Section 4, we report empirical evidence testing different hypotheses of public sector growth for government expenditures in Austria since the mid-1950s. Finally, Section 5 concludes.

2. Descriptive Empirical Evidence

To obtain a picture of the development of the public sector, we concentrate on ratios of government expenditures to gross domestic product. This is superior to a ratio of government revenues to GDP, because expenditures (including those financed by public borrowing) give a better indication of the amount of economic resources absorbed by and allocated through the public sector. What belongs to the government sector, and hence which items should be included in government expenditures, is a question where different answers that are often not clear-cut are possible. Ideally, for the purpose of our study, a very broad measure should be used, including all levels of the public sector (the central government, local governments and municipalities, as well as off-budget public activities and near-public decision-units) and all kinds of public expenditures (consumption, investment, transfers, etc.). Unfortunately, reliable data about relatively broad aggregates are available in Austria only since 1954.[i] For earlier periods, much narrower figures have to be used, namely a series of public consumption (available from 1924 to 1937 and since 1948), and a series of expenditures in the central government's budget (available from 1868 to 1937 and since 1946). The latter series has the disadvantage of referring to a much larger territory, the Austrian part of the Habsburg Monarchy, until 1918. Furthermore, from 1938 to 1945 Austria was occupied by Germany; hence "Austrian" budgets did not exist at all.

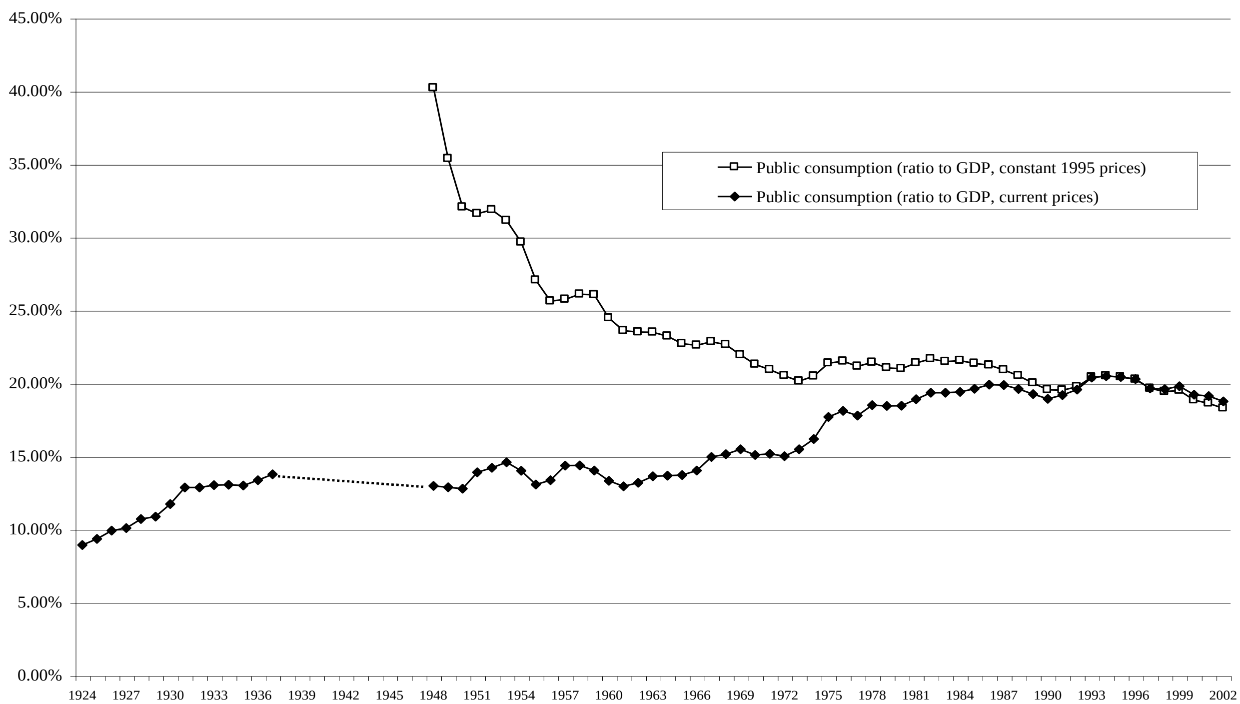
Comparable data for government expenditures and revenues are available on an annual basis since 1870. They refer to the "Austrian" half of the Austro-Hungarian (Habsburg) Monarchy, also called Cisleithania. Data of the Austrian (Cisleithanian) central-state budget (actual expenditures as given in the "Centralrechnungsabschluss") may be regarded as indicators of government expenditures for the period 1870 to 1913. The size of the government sector can then be measured by the ratio of central-state expenditures to GDP for the Austrian part of the Habsburg Monarchy. The development of this series is displayed in Figure 1. Although these figures must be regarded as tentative, they nevertheless show a clear tendency of growth of the government sector, both in absolute and in relative terms, already before World War I. As has been shown by Wysocki (1975), this growth was mainly due to the expansion of state expenditures for infrastructure in a wider sense, including transport and telecommunication, education and science, administration and jurisdiction. In particular, the share of expenditures for railways in total state expenditures grew considerably.

Figure 1: Development Of Central State (Federal Government) Expenditures, 1870-2002
(Government Expenditure-To-GDP Ratio, In %, Nominal)



No data on GDP is available during World War I, and no meaningful estimates can be obtained for nominal GDP during the years of hyperinflation after the war, so the series can be resumed from 1924 only, now for the territory of the Republic of Austria. For the years 1924 to 1937, National Income Account data were estimated by Austrian statisticians. Although their basis is much less reliable than that of the data for the Second Republic, they can at least provide a basis for assessing the development of general economic activity (by real and nominal GDP) and of the government's share of it (by public consumption, both real and nominal). Of course, the latter are quite different from those obtained from the federal budget, which also includes expenditures on investment, transfers, and other items, but excludes lower level expenditures. The ratio of federal budget expenditures to GDP is shown in Figure 1, the ratio of public consumption to GDP in Figure 2.

Figure 2: Development Of Public Consumption, 1924-2002
(Public Consumption-To-GDP Ratio, In %, Nominal And Real, Based On 1995 Prices)

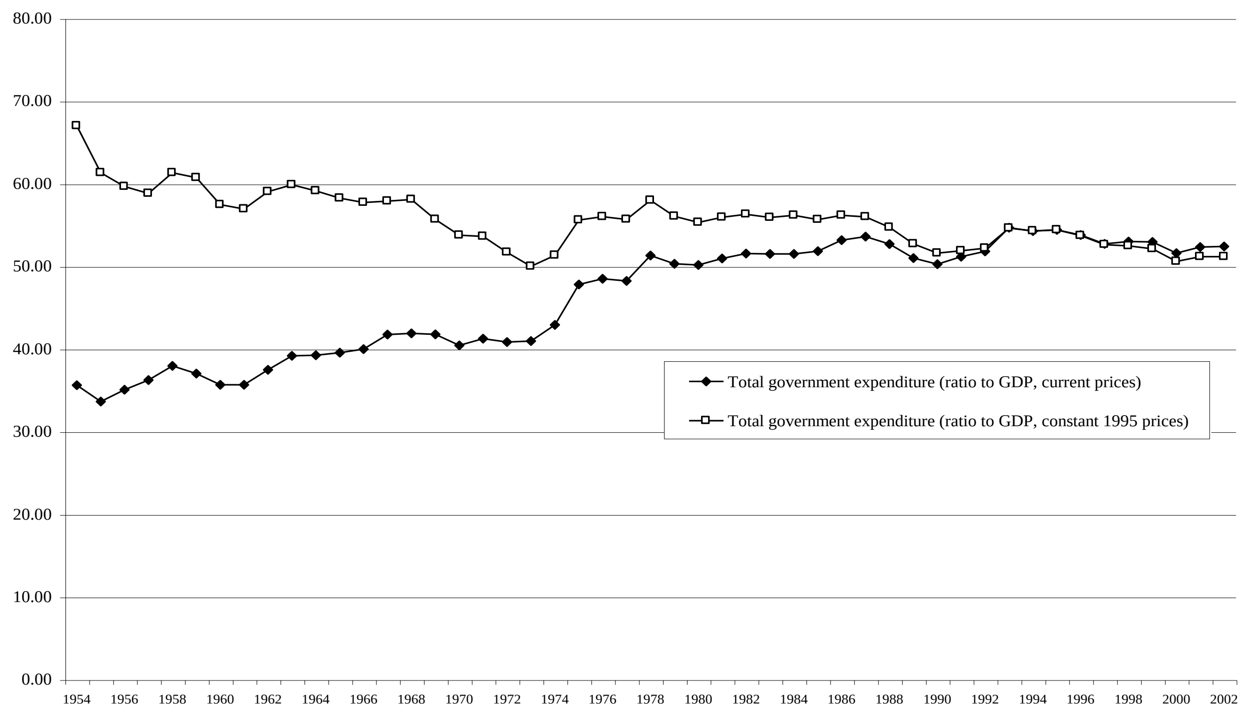


In the years immediately after World War I, state intervention grew due to the public organization of the provision for basic needs, the introduction of laws improving social security, especially unemployment benefits, and of subsidies for food. Government expenditures for the inherited over-sized bureaucracy and the repayment of war loans created additional pressure on the federal budget. Following a hyperinflation from 1921 to 1923, stabilization was brought about and inflation ceased very quickly. After five years of fairly rapid economic growth, the Great Depression hit Austria in a particularly severe way in 1929, with high unemployment in the 1930s, leading eventually to Austria being taken over by the Nazis in 1938. During this period, public consumption showed a more stable development than GDP, hence its share was rising during the depression and stagnation period. On the other hand, the overall federal budget data exhibit considerable fluctuations: the government share, as measured by the ratio of federal budget expenditures to GDP, rose first during the depression and then declined during the long stagnation period.

Austria regained its independence after World War II in 1945. The overall development of the Austrian economy in the Second Republic may be classified into three periods: a period of reconstruction from 1945 to the beginning of the 1950s; a period of rapid growth until the first oil-price shock of 1973/74; and a period of slower economic growth since then. Figures 1 and 2 also show the development of the previously discussed measures of public sector size for the Second Republic. In addition, since 1954 more detailed data from the National Income Accounts are available from Statistik Austria, the Austrian statistical agency. The development of the ratio of total government expenditures to GDP, which is one of the broadest measures of government size for Austria, is displayed in Figure 3. Total government expenditures include public consumption, public investment, transfers and subsidies, and interest payments for government debt.

During the years after the war, the public sector was extended by government investment programs in infrastructure and by the expansion of the system of social security. Financing these expenditures did not provide major difficulties during the years of rapid economic growth, but turned out to be much more problematic since the end of that period. The first oil-price shock in 1973 and its effects on the international economy are the major causes for the recession, which started late in 1974 in Austria. This introduced a period of reduced average growth of the Austrian economy. Stabilization policies first reacted on these developments by a combination of discretionary measures and institutional arrangements that later has been dubbed "Austro-Keynesianism". Full employment was considered to be the primary policy goal, and highly expansionary federal budgets were enacted to counteract the unfavorable influences from the world economy. Negative consequences of these fiscal policies were high budget deficits and the resulting rise of public debt. Another effect of these developments is the rise of the government share in the economy since 1975, which is reflected particularly in the ratio of federal budget expenditures to GDP (Figure 1).

Figure 3: Development Of Total Government Expenditures, 1954-2002
(Total Government Expenditure-To-GDP Ratio, In %, Nominal And Real)



Looking at the development and the composition of total government expenditures shows that there was some growth of nominal government expenditures (its ratio to nominal GDP, see Figure 3), but this did not affect all of its components in the same way. In the period following the first oil price shock, a shift occurred from the "classical" activities of the government sector, such as administration, jurisdiction, and provision of infrastructure, to taking over social and economic risks. The shares of transfers and particularly of interest payments in total government expenditures have grown, those of public consumption and especially of public investment have fallen. Among transfer payments, those for social security have grown particularly fast, especially those for pensions from compulsory pension insurance. Two factors are responsible for this development: first, the number of pensions has increased considerably, because pensions were granted to more and more groups, there were more preliminary retirements, and life expectancy increased; second, per capita payments have been extended, not only in line with economic developments, but also by discretionary measures.

Public fixed investment during the period of rapid economic growth rose faster than GDP (both in nominal and real terms), while in the period since 1975, not only its share in GDP but also its absolute values (even in nominal terms) have fallen. Part of this effect is due to public investment expenditures being increasingly delegated to off-budget agencies; due to lack of reliable information on this, we have to neglect these off-budget expenditures. Public consumption expenditures, on the other hand, accelerated its growth in nominal terms in the late 1970s, but since the 1980s, its shares (both nominal and real) in GDP remained approximately constant. This is mostly due to price effects, because the public consumption deflator grew much faster than the GDP deflator, especially in the 1950s and 1960s, while this relative price effect is no longer present since the 1980s. The main component of public consumption expenditures, outlays for public employees, has shown more than proportional growth in the 1970s because public sector employment has risen faster than total employment and per capita incomes of public employees have grown slightly faster than average incomes in the entire economy. The main contributors to the growth of the public sector over the period since 1954 have been transfers (including subsidies) and interest payments for public debt.

3. Explanations Of Public Expenditure Growth And Their Tests

Towards the end of the 19^{th} century, German economist and public finance professor Adolph Wagner (1835–1917) stated a “Law of Increasing Government Activity” (“Gesetz der wachsenden Ausdehnung der öffentlichen und speciell der Staatsthätigkeiten”) and tried to corroborate this “law” by means of empirical analysis. He concluded that the public sector was growing both in relative and in absolute terms within a national economy, and that this expansion was a burden for the private sector of the economy. Later on, the scientific debate on “Wagner’s Law” has led to several extensions and modifications of the hypothesis and to various approaches to explore its empirical validity.

As an extension of “Wagner’s Law” and alternative explanation of the increase in the level of government expenditures, Peacock and Wiseman (1961) developed the theory of the “Displacement Effect”. It assumes that government expenditures do not grow constantly, but that they grow with “structural breaks”, e.g. wars or important ideological changes in government policies. The increase in the level of expenditures following such an event is not reduced after the event; instead, government expenditures remain at the higher level.

The international literature on the growth of the public sector has not yet developed a generally accepted theory explaining the development of government expenditures. Most work in this field starts from “Wagner’s Law”, adds (more or less ad-hoc) other potential explanatory (groups of) variables, and tries to validate the resulting hypotheses. While until the 1980, estimations of reduced form models were the main approach to test “Wagner’s Law” and related theories of government expenditures’ growth, more modern approaches use time series models and tests for stationarity of different expenditure ratios as well as the exploration of long-term cointegration relations between the expenditure ratio, income and relative prices. The first group of approaches, the estimation of reduced form (mostly: single equation) models, rests on variants of the following equation:

$$G = f(Y, POP, RP) , \tag{1}$$

where G denotes (nominal or real) government expenditures (total government expenditures or only a part of them, such as public consumption), Y is (nominal or real) income (GDP), and *POP* denotes population. A variant of this basic model uses ratios (g as the nominal or real ratio of G to the respective GDP) instead of levels. Sometimes real GDP per capita (*YR/POP*) is used as explanatory variable. An important extension consists in including relative prices *RP*, namely the ratio of the price level of public goods to that of private goods (in effect, a ratio of the deflator for public consumption or total government expenditures to the GDP deflator or the deflator of expenditures of private households).

The theoretical expectations of the arguments in equation (1) are as follows. Regarding the income variable, "Wagner's Law" is confirmed if the income elasticity of the demand for public goods is higher than that for private goods. In this case, higher income (Y) ceteris paribus leads to higher demand for public goods (G). This approach is one of the oldest methods to explore "Wagner's Law" empirically. Regarding the explanatory variable *POP* (population), the direction of the influence on government expenditures is unclear from a theoretical viewpoint. A negative effect may be expected if increasing economies of scale reduce the per-capita costs of public goods. A positive effect may be expected if increases of government expenditures are not used for the production of public goods but for transfer payments, as opportunities for free-riding are more extensive within a larger population. Finally, population growth may have external effects and thus a positive influence on the size of the public sector.

A crucial determinant of rising government expenditures may be the relative prices of public goods as compared to private goods (variable *RP* in equation (1)). The hypothesis of "Baumol's Cost Disease" assumes that labor productivity grows below average in the public sector (particularly regarding services), with a parallel development of wages. This means that expenditures for public services increase without an adequate growth of output. The "Cost Disease" hypothesis regards the amount of public goods as being primarily determined on the supply side: an increase of the price level for public goods leads ceteris paribus to growth of the public sector if the demand for publicly provided goods and services is inelastic with respect to their prices and elastic with respect to income.

Relation (1) has been tested in various ways. Peacock and Scott (2000) give an overview of the specifications and the econometric methods employed during the past decades. In recent years, the rise of "theory-free" econometric methodologies for exploring the characteristics of time series has led to a revival of tests of "Wagner's Law". Such tests start by examining the characteristics and the development of a variable over time, e.g. the nominal government expenditure-to-GDP ratio. The aim of such an exploration is to test for stationarity or the existence of systematic growth trends (deterministic or stochastic) for the time series under consideration. Next, similar methods are used for testing potential cointegration relations between two or more time series. In our case, it can be tested to what extent the time series of equation (1) develop in a parallel manner and whether causal relations exist between them. The aim consists in avoiding "spurious" correlation in applying OLS estimation procedures and detecting long-run relations between possibly non-stationary time series.

Another group of theoretical explanations of government growth comes from the public-choice literature. One of these hypotheses is that of fiscal illusion. This theory suggests that reduced voters' awareness of the burden of financing government expenditures results in an increase of the government expenditure ratio. The rationale for such a theory comes from theoretical models of the behavior of the key actors in the political system (among others, "Leviathan" models). The burden of government growth can be concealed, for instance, by increasing the complexity of the tax system or the extent of credit financing of public expenditures. In empirical studies, the complexity of the tax system is often operationalized by the Herfindahl-index of the concentration of different tax and revenue categories (e.g., income tax, indirect taxes, etc.); a higher index expresses a less complex and thus more transparent tax system. According to the theory of fiscal illusion, a negative coefficient for this variable can be expected. In order to operationalize the extent of credit financing of public expenditures, variables related to the fiscal budget deficit are used. The expected sign of these variables is positive.

Other public-choice explanations of government growth relate the phenomenon to ideologies and party competition or to election cycles. For example, it is conjectured that "left-wing" parties (such as Social Democrats in Europe or Democrats in the US) are more interested in expanding the public sector because they believe in its ability to provide a more "fair" distribution or because their voters rely more heavily on transfers than those of "right-wing" parties (Liberals and Conservatives in Europe, Republicans in the US). In theories of the election cycle, it is conjectured that incumbent politicians tend to "buy" votes by raising transfers and other popular government expenditures just before general elections. All these theories are oriented towards explaining short-run changes in the size of the public sector instead of its long-run growth. If the latter is to be explained, these theories have to be combined with elements of other hypotheses such as the "Displacement Effect" (for example, to explain why higher government expenditures remain after the election or after the "left-wing" parties' stay in office).

4. Econometric Analysis Of Government Expenditures In Austria

4.1. Stationarity And Cointegration

As can be seen from Figure 3, the total government expenditure ratio (ratio of total nominal government expenditures to nominal GDP, both from the Austrian national income accounts) grew on average over the period 1954 to 2002. We want to clarify whether this growth process follows a stochastic trend, i.e. whether the time series is stationary or follows a random-walk process (possibly with a drift and a deterministic trend). In order to test for stationarity, the Augmented Dickey-Fuller test (ADF test) is applied to test for the existence of a unit root. If the time series in levels is non-stationary, we will further test whether the first differences of the series are stationary.

Empirically, the ADF test includes an OLS estimation of the following equation (TGY_t denotes the time series of the nominal total government expenditure ratio):

$$(TGY_t - TGY_{t-1}) = [\beta_0] + \beta_1 TGY_{t-1} + \beta_2 (TGY_{t-1} - TGY_{t-2}) + [\beta_3 Trend], \qquad (2)$$

where β_i, i = 1,2,3, are the coefficients to be estimated. The brackets […] indicate that these terms (the constant and the trend) can be omitted. If the time series has a unit root and is hence non-stationary, the coefficient of the lagged endogenous variable is not significantly different from zero. If the estimated coefficient is significantly negative, the null hypothesis that the time series has a unit root and hence is non-stationary has to be rejected. The ADF test statistic corresponds to the t-statistic for the coefficient of the lagged first difference of the tested time series in equation (2) but does not follow the standard t-distribution. If the time series is identified as non-stationary in levels, the test is repeated for the first difference, i.e. the time series to be tested is (TGY_t – TGY_{t-1}) instead of TGY_t. The equation to be estimated is then

$$\begin{aligned} &((TGY_t - TGY_{t-1}) - (TGY_{t-1} - TGY_{t-2})) = \\ &= [\beta_0] + \beta_1 (TGY_{t-1} - TGY_{t-2}) + \\ &+ \beta_2 ((TGY_{t-1} - TGY_{t-2}) - (TGY_{t-2} - TGY_{t-3})) + [\beta_3 Trend] \end{aligned} \qquad (3)$$

Besides the ADF test, the Phillips-Perron test (PP Test) is an alternative econometric device frequently applied to test for stationarity of a time series.

Table 1 shows the results of estimating equation (2) for the nominal ratio of total government expenditures to GDP (variable TGY_t). The test for stationarity of the time series in levels results in an insignificant ADF test statistic of –1.7054 (compared to the critical value of –3.1828 at the 10% level of significance) with constant and trend and of –1.798 without the trend (Est. 1 and 1a). The estimated equation shows that the time series seems to follow a stochastic but not a deterministic trend (indicated by the insignificant trend variable and the weakly significant constant). The PP test confirms this result with a PP test statistic of –1.4603 compared to a critical value of –3.1816 at the 10% level of significance. While the time series seems to be non-stationary in levels, the ADF test indicates stationarity for the first difference of the time series (Est. 2). This result indicates that the total government expenditure ratio is integrated of order one (I(1)) in the period 1956–2002.

As discussed in Section 2 and shown in Figure 3, there is a substantial difference between the nominal and the real ratio of total government expenditures to GDP. The different development of the price levels in the public and private sector is an indicator of "Baumol's Cost Disease". It leads to fundamentally different processes over time: while non-stationarity in levels can be confirmed for the nominal government expenditure ratio, the real government expenditure ratio has to be judged differently. Figure 3 shows a decrease of the real government expenditure ratio at the beginning of the period and then a rather constant development. This inspection of the graphical representation indicates stationarity of the time series of the real ratio of total government expenditures to GDP, which is confirmed by the respective ADF and PP tests. The ADF test statistic is –3.241, which is smaller than the 10% critical value of –3.1828. A similar result is obtained by applying the PP test, leading to a test statistic of –4.409 which is smaller than the critical value of –3.5045 at the 5% level of significance. These results indicate that the time series of the nominal ratio of total government expenditures to GDP is an I(1)-variable and hence non-stationary in levels, but the real total government expenditure ratio more likely represents a stationary time series.

Table 1: Stationarity Test Of The Nominal Government Expenditure-To-GDP Ratio Time Series Tgy_t And Its First Difference

Tested variable Dependent variable of ADF-test	TGY_t $(TGY_t - TGY_{t-1})$	TGY_t $(TGY_t - TGY_{t-1})$	$(TGY_t - TGY_{t-1})$ $((TGY_t - TGY_{t-1}) - (TGY_{t-1} - TGY_{t-2}))$
	Est. 1 Coefficient (t-statistic)	Est. 1a Coefficient (t-statistic)	Est. 2 Coefficient (t-statistic)
Constant	5.1529 (2.0551*)	2.5969 (2.0596*)	0.3543 (1.7728(*))
TGY_{t-1}	–0.1393 (–1.7054)	–0.0484 (–1.7980)	
$(TGY_{t-1} - TGY_{t-2})$	0.1926 (–1.2900)	0.1274 (0.9151)	–0.9668 (–5.0236**)
$((TGY_{t-1} - TGY_{t-2}) - (TGY_{t-2} - TGY_{t-3}))$			0.1560 (1.0876)
Trend	0.0478 (1.1776)		
ADF-test statistic	–1.7054	–1.7980	–5.0236**
ADF critical value (1%)	–4.1630	–3.5745	–3.5778
ADF critical value (5%)	–3.5066	–2.9241	–2.9256
ADF critical value (10%)	–3.1828	–2.5997	–2.6005
Adj. R^2	0.0508	0.0425	0.4087
S.E. of the regression	1.2316	1.2370	1.2618
F-Statistic	1.8209	2.0202	16.5510**
Breusch-Godfrey LM-Test	1.0101	1.5456	1.4133
Log-Likelihood	–74.3893	–75.1352	–74.4173
Akaike Info. Crit.	3.3357	3.3249	3.3660
Schwarz Crit.	3.4932	3.4430	3.4852
Observations	47	47	46
Period	1956-2002	1956-2002	1957-2002

OLS-estimation; ** p<0.01, * p<0.05, (*) p<0.1

Next, we need to explore whether the ratio of price levels in the private and public sector (variable RP_t) during the same period follows a stationary process. Table 2 shows the estimation results for RP_t. Est. 3 exhibits an insignificant ADF test statistic of –2.4396 (the 10% critical value is –2.5997). This result indicates that the relative price RP_t is a non-stationary time series. From testing the first difference of the time series it can be concluded that RP_t is a I(1)-variable: Table 2 shows in Est. 4 that the 1 % critical level of the ADF test statistic of –3.5778 is larger than the empirically estimated ADF test statistic of –4.2661. This result confirms the visual impression of the development of the time series RP_t. While the development of price levels in the public and the private sector were rather different at the beginning of the period – which may indicate the existence of "Baumol's Cost Disease" –, the differences are much smaller at the end of the period. It can not be excluded that different methods of calculating the deflators, particularly for public consumption, might be a major cause for this. Unfortunately, due to lack of other data and relevant statistical information, such a possibility can not be examined further.

Table 2: Stationarity Test Of The Relative Price Rp_t
(Public-To-Private Sector Price Index Ratio) And Its First Difference

Tested variable Dependent variable of ADF-test	RP_t $(RP_t – RP_{t-1})$	$(RP_t – RP_{t-1})$ $((RP_t – RP_{t-1}) – (RP_{t-1} – RP_{t-2}))$
	Est. 3 Coefficient (t-statistic)	Est. 4 Coefficient (t-statistic)
Constant	2.8712 (3.3832**)	0.7869 (3.0050**)
$RP_{t–1}$	–0.0228 (–2.4396)	
$(RP_{t–1} – RP_{t–2})$	0.1384 (0.9499)	–0.7153 (–4.2661**)
$((RP_{t-1} – RP_{t-2}) – (RP_{t-2} – RP_{t-3}))$		–0.1139 (–0.8309)
ADF-test statistic	–2.4396	–4.2661**
ADF critical value (1%)	–3.5745	–3.5778
ADF critical value (5%)	–2.9241	–2.9256
ADF critical value (10%)	–2.5997	–2.6005
Adj. R²	0.1459	0.4343
S.E. of the regression	1.0686	1.0373
F-Statistic	4.9278*	18.2724**
Breusch-Godfrey LM-Test	4.0972	3.8615
Log-Likelihood	–68.2603	–65.4040
Akaike Info. Crit.	3.0324	2.9741
Schwarz Crit.	3.1504	3.0933
Observations	47	46
Period	1956-2002	1957-2002

OLS-estimation; ** p<0.01, * p<0.05, (*) p<0.1

Next, we test the other potentially influential variable of equation (1), $YRPOP_t$ (real GDP per capita), for stationarity and its order of integration. This variable is important for testing "Wagner's Law. Table 3 shows the results of the ADF test for stationarity of the time series of real GDP per capita ($YRPOP_t$). Based on the results of Est. 5 in that table, the null hypothesis of non-stationarity of the time series is rejected at the 10 % significance level when including a deterministic trend, but not at the 5 % level (ADF test statistic of –3.3629 compared to the 5%-critical level of –3.4952); hence the ADF test is inconclusive. However, when the deterministic trend is excluded from the estimation (Est. 5a), the ADF test statistic is 0.2049, which is much larger than the critical 10%-value of –2.5958. A similar result can be achieved by applying the PP test resulting in a PP test statistic of –2.9716 compared to a critical value of –3.1753. Hence we regard real GDP per capita as non-stationary in levels. Est. 6 shows that the time series of real GDP per capita is stationary in its first difference and can therefore be regarded as an I(1) variable.

As these tests show, all variables examined so far (TGY_t, RP_t, $YRPOP_t$) can be regarded as I(1)-variables. We next examine whether there are cointegration relations between these variables. Bilateral tests for cointegration explore whether a linear combination of two non-stationary time series is itself stationary, i.e. whether the two time series "run parallel" and a long-run relation exists between them. The cointegrating regression for the current problem testing for cointegration between $YRPOP_t$ and TGY_t is

$$TGY_t = \alpha + \beta \cdot YRPOP_t + \varepsilon_t, \tag{4}$$

where α and βdenote the coefficients to be estimated and ε_t is the error term. An OLS estimation of equation (4) delivers two potential indications for cointegration: First, the estimation results in a Durbin-Watson (DW) statistic which can be compared to a critical DW value. The null hypothesis of no cointegration is rejected if the empirical DW test statistic is larger than the critical value. Second, the residuals e_t stemming from the estimation of equation (4) are tested for stationarity (e.g. by means of the ADF test). Stationary residuals indicate cointegrated time series. An OLS estimation of equation (4) leads to superconsistent estimators when the time series are cointegrated. If cointegration cannot be confirmed, an OLS estimation of equation (4) gives spurious results.

Table 3: Stationarity Test Of Real GDP Per Capita ($Yrpop_t$) And Its First Difference

Tested variable Dependent variable of ADF-test	$YRPOP_t$ $(YRPOP_t - YRPOP_{t-1})$	$YRPOP_t$ $(YRPOP_t - YRPOP_{t-1})$	$(YRPOP_t - YRPOP_{t-1})$ $((YRPOP_t - YRPOP_{t-1}) - (YRPOP_{t-1} - YRPOP_{t-2}))$
	Est. 5 Coefficient (t-statistic)	Est. 5a Coefficient (t-statistic)	Est. 6 Coefficient (t-statistic)
Constant	0.8256 (4.9019**)	0.3399 (3.5240**)	0.4083 (4.7889**)
$YRPOP_{t-1}$	–0.3295 (–3.3629(*))	0.0011 (0.2049)	
$(YRPOP_{t-1} - YRPOP_{t-2})$	0.2492 (1.7993(*))	0.0795 (0.5606)	–1.0564 (–5.3478**)
$((YRPOP_{t-1} - YRPOP_{t-2}) - (YRPOP_{t-2} - YRPOP_{t-3}))$			0.1368 (0.9505)
Trend	0.1302 (3.3786**)		
ADF-test statistic	–3.3629(*)	0.2049	–5.3478**
ADF critical value (1%)	–4.1383	–3.5572	–3.5598
ADF critical value (5%)	–3.4952	–2.9167	–2.9178
ADF critical value (10%)	–3.1762	–2.5958	–2.5964
Adj. R^2	0.1457	–0.0322	0.4453
S.E. of the regression	0.2171	0.2387	0.2386
F-Statistic	3.9570*	0.1886	21.4743**
Breusch-Godfrey LM-Test	2.2021	1.1479	0.125033
Log-Likelihood	7.8216	2.2721	2.2749
Akaike Info. Crit.	–0.1442	0.0275	0.0279
Schwarz Crit.	0.0045	0.1390	0.1405
Observations	53	53	52
Period	1950-2002	1950-2002	1951-2002

OLS-estimation; ** p<0.01, * p<0.05, (*) p<0.1

Table 4 shows two estimations of bilateral cointegrating regressions. TGY_t is the dependent variable, while variables $YRPOP_t$ and RP_t are alternatively included as independent variables. The critical DW test statistic of Est. 7 (test for cointegration of TGY_t and $YRPOP_t$) amounts to 0.3453. This value is smaller than the critical value of 0.5110 for the 1% level of significance. The null hypothesis of no cointegration cannot be rejected. A different result is obtained by testing the residuals of Est. 7 for stationarity. A corresponding ADF test leads to an ADF test statistic of –2.1883, which is smaller than the critical value of –1.9478 at the 5% level of significance, but larger than the critical value of –2.612 at the 1% significance level. The hypothesis of no cointegration thus seems to be (weakly) rejected. Note that the coefficient of real GDP per capita is positive as expected from "Wagner's Law".

Est. 8 in Table 4 indicates cointegration of the variables TGY_t and RP_t. The DW value of 0.5154 is larger than the critical value of 0.5110 which indicates that the null hypothesis of no cointegration can be rejected with a probability of 99%. A similar result can be derived by the applying the ADF stationarity test of the residuals: The ADF test statistic of –3.1266 is substantially smaller than the critical value of –2.612 at the 1% level of significance.

Table 4: Estimation Of The Bilateral Cointegration Equations Of The Tgy_t Time Series With Rp_t And $Yrpop_t$

Dependent variable of the cointegrating regression	TGY_t	TGY_t	TGY_t
	Est. 7 Coefficient (t-statistic)	Est. 8 Coefficient (t-statistic)	Est. 8a Coefficient (t-statistic)
Constant	29.2702 (28.4196**)	17.9887 (16.7379**)	11.6886 (5.6766**)
$YRPOP_t$	1.1262 (17.5613**)		–0.7304 (–3.4660**)
RP_t		0.3539 (26.9679**)	0.5706 (8.9658**)
Adj. R²	0.8649	0.9380	0.9498
S.E. of the regression	2.5218	1.7085	1.5378
F-Statistic	308.3992**	727.2654**	454.8451**
Durbin-Watson statistic	0.3453	0.5154	0.5509
Log-Likelihood	–113.8307	–94.7532	–89.0686
Akaike Info. Crit.	4.7278	3.9491	3.7579
Schwarz Crit.	4.8050	4.0263	3.8737
Observations	49	49	49
Period	1954-2002	1954-2002	1954-2002

OLS-estimation; ** p<0.01, * p<0.05, (*) p<0.1

Tests regarding cointegration of the variables $YRPOP_t$ and RP_t give the opposite result. The DW test statistic of 0.1375 is smaller than the critical value even for the 10% level of significance, and a test for stationarity of the residuals results in an insignificant ADF test statistic of –0.5065 (compared to the critical 10% value of –2.5997). As expected a priori, these two variables are not cointegrated with each others.

In contrast to the bilateral tests for cointegration, multilateral cointegration tests explore cointegrating relations between more than two time series simultaneously in vector-autoregressive (VAR) models. The methodology applied here is due to Johansen and Juselius (1990). A VAR system for the variables examined here, namely government expenditure ratio, relative price and real GDP per capita (TGY_t, RP_t, $YRPOP_t$), can be specified in the following way (including a constant but not a deterministic trend):

$$TGY_t = \alpha_{10} + \sum_{j=1}^{p} \beta_{11j} TGY_{t-j} + \sum_{j=1}^{p} \beta_{12j} RP_{t-j} + \sum_{j=1}^{p} \beta_{13j} YRPOP_{t-j} + \varepsilon_{1t}, \tag{5}$$

$$RP_t = \alpha_{20} + \sum_{j=1}^{p} \beta_{21j} TGY_{t-j} + \sum_{j=1}^{p} \beta_{22j} RP_{t-j} + \sum_{j=1}^{p} \beta_{23j} YRPOP_{t-j} + \varepsilon_{2t} \tag{6}$$

$$YRPOP_t = \alpha_{30} + \sum_{j=1}^{p} \beta_{31j} TGY_{t-j} + \sum_{j=1}^{p} \beta_{32j} RP_{t-j} + \sum_{j=1}^{p} \beta_{33j} YRPOP_{t-j} + \varepsilon_{3t}. \tag{7}$$

The aim of the testing procedure is to determine the number of cointegrating relations between the time series of the VAR system. For this purpose, the equations are rewritten for the first difference of the dependent variables and transformed accordingly. The multilateral Johansen-Juselius cointegration test explores different hypotheses, depending on the existence of a deterministic trend (linear or quadratic) and/or a constant in the equations. The respective null hypotheses assume a certain number of cointegrating relations. The number of possible cointegrating relations can be between zero and *k*, where *k* is the number of time series in the system. In the current context, the minimum number of cointegrating relations is zero (no cointegration), the maximum is 3 (which would imply stationarity of all three series).

Table 5: Estimation Of Multilateral Cointegration Relationships In A VAR System Of The Tgy_t, Rp_t, $Yrpop_t$ Time Series (Johansen-Juselius Procedure)

	Likelihood ratio	Critical value (5%-level)	Critical value (1%-level)	No. of cointegrating relations
Lags: 2, Assumptions for the test of H_0: Time series without deterministic trend, cointegrating regression without constant				
	37.9735	24.31	29.75	None **
	7.8382	12.53	16.31	At most 1
	2.7250	3.84	6.51	At most 2
Lags: 2, Assumptions for the test of H_0: Time series without deterministic trend, cointegrating regression with constant				
	45.6943	34.91	41.07	None **
	9.1507	19.96	24.60	At most 1
	4.0337	9.24	12.97	At most 2
Lags: 2, Assumptions for the test of H_0: Time series with linear trend, cointegrating regression with constant				
	20.0031	29.68	35.65	None
	5.2971	15.41	20.04	At most 1
	0.3015	3.76	6.65	At most 2
Lags: 2, Assumptions for the test of H_0: Time series and cointegrating regression with linear trend				
	29.6226	42.44	48.45	None
	14.6274	25.32	30.45	At most 1
	4.1676	12.25	16.26	At most 2
Lags: 2, Assumptions for the test of H_0: Time series with quadratic trend, cointegrating regression with linear trend				
	24.5834	34.55	40.49	None
	10.0081	18.17	23.46	At most 1
	0.0307	3.74	6.40	At most 2

** $p<0.01$, * $p<0.05$ for the rejection of the null hypothesis that there are *k* cointegrating relations between the time series.

Table 5 shows the results of the empirical tests regarding the number of cointegrating relations in the VAR system (5)–(7). As maximum lag length, we take 2. For each of the five possible assumptions, the null hypotheses of k cointegration relations are tested against the alternative of k+1 cointegration relations. Under two assumptions (each without a deterministic trend), the null hypothesis of no cointegrating relation has to be rejected at the 1% level of significance. From this, it seems as if the system includes at least one cointegrating relation. Under the three other assumptions tested, the null hypothesis of no cointegrating relation cannot be rejected. Applying the Johansen-Juselius test procedure is thus inconclusive: the system exhibits either a deterministic trend or a (probably only one) cointegrating relation. This result does not corroborate the results obtained by the bilateral cointegration tests presented above which pointed towards cointegrating relations between TGY_t and RP_t, and between TGY_t and $YRPOP_t$, but not between RP_t and $YRPOP_t$.

4.2. Structural Econometric Explanations For The Growth Of Government Expenditures

The results obtained so far are problematic for further analysis. It seems relatively clear that all time series considered are non-stationary in levels but stationary in their first difference, i.e. I(1). It is, however, not obvious whether there exist cointegrating relations between the ratio of government expenditures to GDP (expenditure ratio) on the one hand, and the relative price (ratio of the price levels of public to private goods) and/or GDP per capita on the other hand. If there are cointegrating relations, static equations for long-term relations and error-correction specifications for short-term adjustments between the expenditure ratio and its possible determinants can be estimated by OLS. If there is no cointegration, however, OLS estimations will produce only spurious correlation. In this case, there is no long-run relation between the variables under consideration, and equations in first differences can be estimated as the first differences are all stationary.

In the following, two main hypotheses will be tested for Austria:

- "Wagner's Law": the validity of this hypothesis can be regarded as confirmed under the assumption of cointegration by a significant positive coefficient for the explanatory variable $YRPOP_t$.
- "Baumol's Cost Disease" of the public sector: under cointegration, a positive and significant coefficient for the explanatory variable RP_t is an indication for the validity of the hypothesis, i.e. that the growth of the public sector is at least partially determined by higher price increases of public goods and services .

These two hypotheses are first tested under the assumption of cointegrating relations between the two explanatory variables and the nominal expenditure ratio, using OLS estimations in levels of the respective variables. The results of these estimations are not very encouraging and are not presented in detail here. To be more specific, the estimation of an OLS regression of the expenditure ratio on real GDP per capita and the relative price (Est. 8a in Table 4) yields a significant but negative coefficient for $YRPOP_t$, which is completely against the theoretical expectations from "Wagner's Law". On the other hand, we get a significant positive coefficient for RP_t. The estimated equation shows strong positive serial correlation of the residuals. Trying to correct for that, the estimated autocorrelation coefficient becomes close to one. A unit root of the residuals is thus more likely than the temporarily assumed cointegration. Estimating the equations with the independent variables separately results in coefficients that are very different from those in the regression including both. The coefficient of the income variable now changes the sign and becomes (significantly) positive (see Table 4, Est. 7). Attempts at estimating error-correction models are also not successful.

Therefore, we assume that there are no cointegrating relations between the variables and estimate the following equations in first differences instead of levels. From the perspectives of the underlying (rather rudimentary) theories, such an approach has to be considered as second-best only as both the "Cost Disease" hypothesis and "Wagner's Law" refer to long-term relations, while an estimation in first differences presumes short-term relations only and does not allow for a long-term equilibrium. From an econometric viewpoint, however, such an approach turns out to be successful, as can be seen from inspection of Table 6.

Est. 9 in Table 6 is the counterpart of the aforementioned estimation in levels. Both presumably influential variables are highly significant, and the adjusted R^2 of 0.5416 is high for a regression in first differences. Autocorrelation of the residuals cannot be excluded, but is less of a problem than in the estimation attempts described above. Est. 10 and 11 show that both independent variables remain significant also when included separately in the equation. At first sight, the negative sign of the coefficient for the GDP variable is surprising. It is apparently not compatible with "Wagner's Law", but the "Law" can be considered as being falsified already by the failure to find a robust long-term equilibrium relation when estimating the equation in levels. Only short-term fluctuations can be described by a relation between changes of variables. Such a relation, however, can be given a plausible interpretation in the current case: due to automatic stabilizers on the expenditure side (e.g., unemployment benefits) and discretionary policy measures (e.g., stabilization policies by means of public purchases), a decline of GDP per capita can result in a rise of the expenditure ratio, ceteris paribus. In periods of economic booms, the relation runs the other way round.

Thus, instead of "Wagner's Law", an influence of the business cycle on government expenditures can be identified. This can be explored also by other variables changing with business cycles. For instance, Est. 12 and 13 in Table 6 show the influence of the rate of unemployment (UR_t, measured by the national "Austrian" approach). The rate of unemployment has in fact a significantly positive influence on the expenditure ratio, as expected. This influence, however, is highly (negatively) correlated with that of GDP per capita, casting doubt on the reliability of Est. 13. A comparison between Est. 12 and 9 shows that the inclusion of the unemployment rate instead of GDP per capita significantly reduces the statistical fit of the estimated equation. A comparison between Est. 12 and Est. 13 shows that the size of the estimated coefficient of the unemployment rate is not stable. The results obtained so far are similar to those from a previous study (Neck and Schneider, 1988): The validity of "Wagner's Law" cannot be confirmed for Austria in the period 1954–2002. The government expenditure ratio does not increase systematically with higher real income per capita, but depends negatively on short-term income fluctuations. The "Cost Disease" hypothesis is validated, i.e. a higher price increase in the public sector as compared to the private sector leads to an increase of the government expenditure ratio. An influence of the business cycle can be shown, as the expenditure ratio increases with a decline of real GDP per capita and (more weakly) with an increase of the unemployment ratio.

Table 6: Explanation Of The Development Of The Government Expenditure-To-GDP Ratio Tgy_t
(All Variables In First Differences)

Dependent Variable	TGY_t	TGY_t	TGY_t	TGY_t	TGY_t
	Est. 9 Coefficient (t-statistic)	Est. 10 Coefficient (t-statistic)	Est. 11 Coefficient (t-statistic)	Est. 12 Coefficient (t-statistic)	Est. 13 Coefficient (t-statistic)
Constant	1.2308 (4.5225**)	1.7001 (5.8986**)	–0.1127 (–0.4275)	–0.2306 (–1.1170)	0.8149 (3.0869**)
$YRPOP_t$	–3.6893 (–6.8011**)	–3.4153 (–5.4504**)			–2.7456 (–5.0867**)
RP_t	0.4624 (4.1595**)		0.3706 (2.3847*)	0.4819 (3.9282**)	0.5071 (5.1441**)
UR_t				1.3932 (5.5531**)	0.8536 (3.7508**)
Adj. R^2	0.5416	0.3792	0.0907	0.4484	0.6448
S.E. of the regression	0.8776	1.0214	1.2361	0.9627	0.7726
F-Statistic	28.7682**	29.7073**	5.6866*	20.1062**	29.4381**
Durbin-Watson statistic	1.3759	1.3769	1.5791	1.4474	1.2300
Log-Likelihood	–60.2956	–68.1033	–77.2635	–64.7373	–53.6371
Akaike Info. Crit.	2.6373	2.9210	3.3026	2.8224	2.4015
Schwarz Crit.	2.7543	2.9989	3.3806	2.9393	2.5575
Observations	48	48	48	48	48
Period	1955-2002	1955-2002	1955-2002	1955-2002	1955-2002

OLS-estimation; ** p<0.01, * p<0.05, (*) p<0.1

As mentioned in Section 3, there are also several politico-economic theories trying to explain the development of government expenditures. Here we concentrate on some approaches that may be relevant for the Austrian situation. The results are presented in Table 7. To test for fiscal illusion influences, the Herfindahl-index of tax concentration $THERF_t$ and the ratio of the net budget deficit of the central government to GDP NSY_t are included as regressors. The results of Est. 14 and 15 in Table 7 show that both explanatory variables are significant with the expected sign. Both also contribute separately to the fit of the OLS estimation in first differences. For explaining a long-term relation, fiscal illusion would be a problematic assumption, because an ever increasing fiscal illusion of voters seems to be rather implausible.

Table 7: Explanation Of The Development Of The Government Expenditure-To-GDP Ratio Tgy_t (All Variables Except Dummies In First Differences)

Dependent Variable	TGY_t	TGY_t	TGY_t	TGY_t	TGY_t	TGY_t
	Est. 14 Coefficient (t-statistic)	Est. 15 Coefficient (t-statistic)	Est. 16 Coefficient (t-statistic)	Est. 17 Coefficient (t-statistic)	Est. 18 Coefficient (t-statistic)	Est. 19 Coefficient (t-statistic)
Constant	0.8818 (4.9329**)	0.9786 (5.8941**)	0.9048 (5.3924**)	0.9681 (5.5754**)	0.9920 (5.1408**)	1.1500 (5.3286**)
$YRPOP_t$	–2.1558 (–5.7573**)	–2.3108 (–6.4058**)	–2.3237 (–6.5881**)	–2.3237 (–6.2995**)	–2.3159 (–6.3159**)	–2.3646 (–6.5460**)
RP_t	0.2473 (3.2686**)	0.2151 (2.9592**)	0.2039 (2.8584**)	0.2176 (2.9299**)	0.2165 (2.9164**)	0.2013 (2.7530**)
$THERF_t$	–78.0165 (–5.8441**)	–85.5510 (–6.9565**)	–84.6192 (–7.0319**)	–85.2479 (–6.8187**)	–85.4306 (–6.8512**)	–84.8854 (–6.9372**)
NSY_t	0.3337 (3.4833**)	0.3569 (3.7451**)	0.3351 (3.5658**)	0.3537 (3.6370**)	0.3545 (3.6239**)	0.3378 (3.5191**)
URQ_t	0.2408 (1.3747)					
SPDOM			0.2806 (1.7319(*))			
ELECT				0.0400 (0.2351)		
$DIST\text{-}ELECT_t$					–0.0104 (–0.1418)	
COAL						–0.2048 (–1.2327)
Adj. R^2	0.8376	0.8343	0.8416	0.8305	0.8304	0.8362
S.E. of the regression	0.5224	0.5277	0.5159	0.5336	0.5339	0.5246
F-Statistic	49.4884	60.1434	50.9521	47.0687	47.0223	49.0000
Durbin-Watson statistic	1.4621	1.5244	1.5618	1.5114	1.5121	1.6150
Log-Likelihood	–33.7340	–34.7904	–33.1348	–34.7588	–34.7789	–33.9374
Akaike Info. Crit.	1.6556	1.6579	1.6306	1.6983	1.6991	1.6641
Schwarz Crit.	1.8895	1.8528	1.8645	1.9322	1.9330	1.8980
Observations	48	48	48	48	48	48
Period	1955-2002	1955-2002	1955-2002	1955-2002	1955-2002	1955-2002

OLS-estimation; ** $p<0.01$, * $p<0.05$, (*) $p<0.1$

Est. 16 in Table 7 presents a specification that tests the hypothesis that the dominance of "left-wing" parties in the government increases the government expenditure ratio more than governments in which the conservative Austrian People's Party (ÖVP) participated. The variable *SPDOM* is a dummy taking the value of one for the period 1971 to 1986 and zero otherwise. For this period of either an absolute majority of the Austrian Social Democratic Party (SPÖ) or a short-term coalition ("small coalition") of the Social Democrats with the Austrian Freedom Party (FPÖ), it can be hypothesized that a social-democratic approach to fiscal policy was prevalent and that Social

Democratic governments were more likely to increase government expenditures than conservative governments. The regression shows the expected positive coefficient which is, however, not significant. While the size and significance of the other variables in the equation does not change, an inclusion of the variable *SPDOM* does not increase the explanatory power of the estimation. An influence of the ideology of the parties in office can therefore not be confirmed. The opposite conclusion that "conservative" governments led by the ÖVP have reduced the expenditure ratio can also not be drawn.

Two more regressions explore the hypothesis of a political business cycle in Austria. The influence of two variables is tested: first, the variable *ELECT* with a value of one for election years (zero otherwise) is tried as additional regressor; second, the variable *DIST* denoting the distance to the next national elections (in years) is included in the estimation. The existence of a political business cycle may lead to higher government expenditures before elections in order to influence the economy positively and to convince voters of the economic expertise of the government, or just to present "voting gifts" to the electorate, while after the elections, a reduction of government expenditures in order to reduce government debt might occur. Est. 17 and 18 in Table 7 show that both variables relating to the political business cycle give the expected sign but are far from being significant. The government expenditure ratio seems to be independent of ideology or the political business cycle. Finally, Est. 19 shows that also the form of the government (variable *COAL* is a dummy for coalition governments) does not provide much explanatory power.

5. Concluding Remarks

The descriptive analysis and the econometric tests applied in this paper confirm the growth of government expenditures, measured as ratio of nominal government expenditures to nominal GDP. If the officially published price indices are correct, there is, however, no growth in real government expenditures (measured as ratio to real GDP). The growth of government expenditures seems thus to be primarily a problem of increasing prices in the public sector and not so much connected to GDP.

An inspection of the time series of the (nominal) government expenditure ratio, GDP per capita and relative prices of public and private goods suggests close relations between these variables. The econometric tests show that all three variables are integrated of order one, i.e. they are non-stationary in levels but stationary in first differences. Further econometric analysis, however, cannot confirm the initially hypothesized cointegration of the time series of government expenditures (as ratio to GDP) on the one hand, and the time series of GDP per capita and relative prices on the other hand. As cointegration is an important precondition for OLS estimations of structural models in levels, regressions were run with variables entering as first differences.

The empirical estimations show that the growth of the nominal government expenditure ratio is strongly influenced by cost increases in the public sector ("Baumol's Cost Disease"), while per-capita income exerts a negative influence. "Wagner's Law" of increases of the government expenditure ratio following increases of per-capita income cannot be confirmed for Austria. Instead, the influence of the income variable can be explained by the influence of short-term fluctuations (the business cycle). For similar reasons, there is a positive influence of the level of unemployment on the government expenditure ratio. A test of the validity of the fiscal illusion hypothesis proves to be successful. The hypothesized influence of other variables related to politico-economic determinants of the government expenditure ratio, such as the ideology of the parties in government, the political business cycle and the form of the government (e.g., coalition governments), does not turn out to be of much explanatory power.

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[i] It is, however, not trivial to construct a coherent time series of government expenditure even for this period since the systems of national income accounting have changed during the last five decades.

NOTES