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by

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CONSUMPTION AND GOVERNMENT DEBT IN HIGH DEFICIT COUNTRIES.
IS TAX-DISCOUNTING STABLE OVER TIME?

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ABSTRACT

In the paper I investigate the extent of tax-discounting effects on private (non-durable and total) consumption in Italy and Belgium, two countries that have experienced dramatic changes in the trends of government debt and deficits in the post-war period. Using cointegration techniques, I estimate both a static (long-run) and a dynamic (error-correction) version of the consumption model. Estimates of the tax-discounting parameter are nil in the long-run relationship and quite large and significant in the dynamic specification, suggesting that tax-discounting behaviour is closely linked to the short-run adjustment components of aggregate consumption behaviour. A stability analysis of the dynamic model, focusing on the parameter of fiscal illusion, suggests that tax-discounting arises only in coincidence with sharp upward movements in the dynamics of public debt. The results are particularly robust for Belgium, where tax-discounting can be successfully modelled as a positive function of the debt-to-GNP ratio. These findings tend to confirm the cross-country evidence provided in Nicoletti (1988b) and suggest that Ricardian phenomena may be contingent on changes to government financing regimes, rather than being a stable characteristic of aggregate consumer behaviour deriving from the nature of individual preferences. When debt dynamics become explosive, a positive relationship between deficits and private saving may result from consumers' expectations of the increased tax burdens needed for a swift return to a stable policy course.

Introduction.

In the last two decades both economic theory and the observation of the behavior of large industrial economies have shaken strong a priori beliefs concerning the impact of government debt financing on private saving. The debate among applied economists has revolved around two main positions, which for simplicity will be called "traditional" and "Ricardian". According to the traditional standpoint government deficits stimulate private consumption and depress private and national saving, crowding out private capital formation in the long-run. On the other hand the Ricardian view considers national saving and economic activity largely independent of the way the government finances its budget: deficit financing generates the private savings necessary to absorb the additional government debt, leaving national saving intact. This view is generally based on the ability of consumers to foresee and discount the future taxes associated with current issues of government debt --a behavior that has been labelled "tax discounting". Both the traditional and the Ricardian positions hinge on assumptions concerning the overall characteristics of the economy: the structure of individual preferences, the constraints faced by agents, the configuration of markets and the nature of government policies. Hence the direction and the magnitude of the effects of government financing policies is largely an empirical question.

Empirical studies dealing with the "tax discounting" issue usually assume that tax discounting (or the lack of it) is a permanent and stable characteristic of consumer behaviour, linked to the underlying structure of preferences. This paper takes a different approach. Its purpose is to show that tax discounting may be driven by contingent factors, such as changes in financing policy regimes of the government. Our conjecture is that the intensity of tax discounting may increase when budgetary policies take an unsustainable course, since in this case consumers cannot expect any longer to see the debt burden shifted beyond their expected lifetimes. In a previous paper by the author [Nicoletti (1988b)], evidence of a relationship between tax-discounting behavior and government financing regimes was presented on a cross-country basis: sizeable tax-discounting effects were detected only in countries whose governments had relied heavily on debt financing over the sample period. In this paper, we concentrate on Italy and Belgium --two countries which have experienced noticeable changes in government financing regimes in the post-war period-- and relate the intensity of tax-discounting to changes in the policy environment over time. First, we try to distinguish between the short-run and long-run impact of government debt on consumption. Applying the unit-root testing strategy proposed by Perron (1988) and the cointegration techniques of Engle and Granger (1987) and Phillips (1988) we derive estimates of the tax-discounting parameter in both static and dynamic (error-correction) versions of the consumption model. Second, we analyse the stability of the tax-discounting parameter over time and estimate a variable-parameter specification of the dynamic model in which changes in government financing policies --as summarized by the dynamics of the debt ratio-- affect the degree of tax discounting by consumers. The analysis is performed for both total and non-durable private consumption over the 1953-87 period.

The main results are the following. First, estimates of the tax-discounting parameter are nil in the static model and quite large and significant in the dynamic specification. An interpretation of this contrasting evidence is that tax-discounting may be linked to short-run, contingent factors --rather than to the long-run, co-trending features of consumption and its determinants. Second, in the dynamic consumption model strong Ricardian effects appear only after sharp upward changes in the trends of government debt and deficits. While in Italy no clear breakpoint can be statistically identified for the tax-discounting parameter, in Belgium a structural break is identified in correspondence to the change in government financing policies and, after the breakpoint, tax-discounting effects can be successfully modelled as a positive function of the debt-to-GNP ratio. This suggests that changes in government debt policies may affect in important ways the relationship between government deficits and private saving. In particular, unstable dynamics of public debt may trigger increases in private saving, which may be explained by the expectation of budget restrictions in the near future.

The plan of the paper is as follows. The next section provides a background discussion that motivates our empirical approach. Section II presents the theoretical consumption model and its empirical implementation, as a cointegrating system and in error-correction form. Section III presents the results of the empirical analysis. First, unit-root tests, cointegration tests and "fully-modified" least squares estimates of the static model are presented. Next, estimates of various versions of the dynamic consumption model are provided and the relationship between tax-discounting and government policies is explored through stability tests and variable-parameter estimates. A concluding section briefly summarizes and interprets the results.

I. Background

A large number of studies has looked at the effects of government debt and deficits on aggregate private consumption[1]. Most of these studies tried to discriminate between the traditional and Ricardian views by fitting to aggregate time-series data two competing microeconomic models of the equilibrium behavior of the consumer: the life-cycle model and the dynastic model[2]. Under certain conditions, these models imply macroeconomic consequences of government deficit financing that are consistent respectively with the traditional and Ricardian views. Thus they apparently make possible to formulate tightly specified econometric models in which the two views can be tested. Unfortunately, after more than a decade of empirical testing the results produced by this research strategy remain contradictory even at the level of single countries analysed over the same sample period.

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[1] This literature is surveyed by Bernheim (1987).

[2] The life-cycle model postulates "selfish" individuals that consume their lifetime resources in order to maximise their own satisfaction during their finite life-spans. The dynastic model postulates "altruistic" individuals that not only choose to consume optimally during their own life-spans but also choose a level of bequests (or gifts) that makes possible the attainment of the maximum welfare by their offsprings (or parents).

There are at least two reasons that could explain the apparent inefficacy of this approach. The first is that the policy implications of the two models are mutually exclusive only under auxiliary assumptions concerning the nature of government policies. Crucial assumptions are that the horizon of government policies is longer than the life-span of the representative consumer and that financing policies are non distortionary and known with certainty. Otherwise, it would be possible to observe crowding-out effects of government deficit policies in a world of dynastic consumers or, alternatively, Ricardian effects in a world of life-cycle consumers[3]. For instance, if debt repayment occurred within the average lifetime of consumers, so that the economic horizon of government financing policies coincided with that of the debt holders, debt issuance could be expected to generate negative income effects independently of the underlying nature of consumer preferences[4]. Moreover, the implied positive effect on saving could be enhanced by precautionary savings, since future expected taxes may increase earnings uncertainty. Therefore in this case, rejection or acceptance of Ricardian behavior at the aggregate level does not necessarily discriminate between equilibrium consumption models.

The second reason is that the relationship between the estimated models and the true data generating process is often doubtful. Generally, the estimated consumption functions impose on the data the straightjacket of tightly specified equilibrium relationships. These specifications have at best oversimplified dynamic features and cannot account for short-run disequilibrium phenomena. This problem is frequently compounded by the use of inappropriate inference procedures. For example, many empirical consumption models are cast in terms of contemporaneous levels of variables, ignoring that an important

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[3]While the first case is widely acknowledged (see for example Barski et al. [1986] and Kimball and Mankiw [1988]), few authors have emphasized the possibility that life-cycle consumers might behave in a Ricardian way. An exception is Chan (1983) who shows that, if the distribution of future taxes across individuals is uncertain, government deficit financing may be associated with incremental private savings caused by precautionary motives.

[4]Poterba and Summers (1987) present estimates for the United States suggesting that only 25 percent of the burden of public debt is shifted on to future generations.

body of evidence suggests that variables in levels are often non stationary over time. Recent work in econometrics shows that standard inference theory cannot be applied in general to estimates of equilibrium models involving non stationary variables. Finally, the stability of the estimated parameters -- including that expressing the degree of tax discounting-- remains usually untested in most of these empirical studies.

Our analysis differs in two ways from this line of research. First, we view the distinction between short-run and long-run behavior as being important for the effects of government debt and deficits on private consumption. These effects depend on agents' expectations about the discounted value of their future stream of incomes, net of "permanent" taxes. Since expectations are revised each period in the light of actual or announced changes in financing policy regimes, an explicit treatment of consumption dynamics is needed. Therefore, we estimate both an unrestricted dynamic error-correction model and its long-run equilibrium solution. The long-run consumption specification is consistent with most equilibrium theories of consumption behavior, while the error-correction representation is a suitable simple representation of the data generating process. At the same time, the traditional and Ricardian views of the effects of debt and deficits on consumption are nested in a simple way in both the long-run and the short-run versions of the model. An ideal bridge between the two specifications is offered by "Granger's representation theorem" [Engle and Granger (1987)] for systems of cointegrated series. Cointegration techniques allow standard inference to be performed in equilibrium models relating non-stationary variables. Moreover, the duality between cointegrated systems and error-correction specifications ensures that both the co-trending characteristics of the variables and their temporary deviations from equilibrium are taken into account. This makes it possible to discriminate between short-run and long-run factors affecting consumption, a desirable feature in trying to estimate the degree of tax discounting by consumers.

Second, we expect the relationship between private consumption and government debt to be characterised by breaks and/or non-linearities over time. Unforeseen changes in financing policy regimes can generate instability in the estimates of the tax-discounting parameter. Moreover, in changing economic conditions, fiscal variables that appeared to be irrelevant for consumption decisions may prove to be important components of the underlying theoretical model. For instance, the size and dynamics of the public debt ratio may affect the degree of tax discounting in private consumption. On the one hand, as public debt grows out of control the horizon of government policies becomes shorter, since in general policies leading to an explosive debt-to-GNP ratio cannot be sustained permanently. Consumers expecting imminent tax increases or reductions in government transfers have good reasons to behave in a Ricardian way[5]. On the other hand, it is reasonable to assume that, when the level of the debt is low, the welfare cost to consumers of ignoring the link between government current debt financing and future taxes is very small, possibly lower than the cost of collecting information about government current and future budgetary policies. But for high enough levels of debt, the proportion of consumers behaving in a Ricardian fashion may become important, because this cost is likely to increase with the value of the outstanding stock of debt[6]. These considerations suggest that the time-series behaviour and the possible determinants of the tax-discounting parameter must be analysed more closely than is usually done.

II. The model

II.1 An equilibrium specification of consumption with Ricardian effects

The basic equilibrium specification of consumption used in the analysis relates real consumption per capita, C, to per capita real disposable income, Y^d, the per capita real stock of financial and non financial wealth, W, the real interest rate, R, and the rate of growth of per capita net real income, Y.

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[5]This is all the more true if the growth of public debt leads to a switch from a regime in which debt can be redeemed through economic growth to one in which government solvency must be ensured by new taxes, a situation likely to occur when interest rates depend positively on the stock of public debt outstanding.
[6]This argument has been formally analysed by Gruen (1988).

Using lower-case letters to indicate logarithms of variables and a hat to indicate first differences, the log-linear version of this relationship is:

$$(1) \qquad c_t = a_0 + a_1 R_t + a_2 w_{t-1} + a_3 y^d{}_t + a_4 \hat{y}_t$$

Equation (1) is consistent with the Life Cycle or Permanent Income hypotheses of consumer behavior once appropriate proxies for the lifetime resources of the consumer are chosen[7]. Imposing the homogeneity restriction $a_2 + a_3 = 1$, the equilibrium average propensity to consume out of disposable income (c/y^d) is constant as long as the real interest rate and the growth rate of income remain unchanged. The dependence of the long-run propensity on the interest rate and on real income dynamics is a standard feature of consumption models derived from consumer optimization in a growing environment[8].

Another important dynamic factor often overlooked by many empirical consumption studies is price inflation. Several theoretical explanations have been suggested for the influence of inflation on real consumption and these have been backed by substantial empirical evidence[9]. In this study we concentrate on the effects due to mismeasurement of real disposable income and loss of purchasing power of financial wealth. Therefore, we adopt a definition

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[7] For similar long-run specifications of consumption see, for instance, Modigliani (1975,1986) and Hendry (1983).

[8] Although these effects are often ignored in consumption function estimates, there are both theoretical and empirical reasons for including them in the equilibrium consumption model. Farrell (1970), Modigliani (1975,1986), Currie (1981) and Nickell (1985) provide theoretical reasons for the presence of interest rate and growth effects; Modigliani (1970), Davidson _et al._ (1978), Feldstein (1980) and Graham (1987) estimate consumption models including growth rates among the regressors.

[9] Theoretical explanations include disequilibrium effects due to misperceptions by agents [Deaton (1977)], effects due to the mismeasurement of real income [Davidson _et al._ (1978)] and effects caused by the erosion of the nominally denominated component of wealth [Jump (1980)]. Supporting empirical evidence was provided by Deaton (1977), Townend (1976), Davidson _et al._ (1978), Hendry and Von Ungern-Sternberg (1981) and Nicoletti (1988b), among others.

of disposable income that nests the conventional National Accounts measure, which includes payments of <u>nominal</u> interests on financial assets held by the private sector, and an alternative measure, called "Hicksian", which includes only payments of <u>real</u> interests[10]. To this purpose a parameter is introduced expressing the degree of Hicksian inflation-correction of disposable income by agents. By assumption this parameter, denoted hc, takes values between zero and unity. It is zero when the conventional measure of disposable income is correct --i.e. agents mistakenly take the inflation premium component of interest transfers as an accrual to real income-- and it is unity when the Hicksian measure is appropriate --i.e. agents correctly treat this inflation premium component as a compensation for the erosion of their stock of real financial wealth.

Nesting the traditional and Ricardian views of the effects of government debt and deficits on consumption requires further departures from the standard definitions of income and wealth. To this purpose, we introduce a parameter, denoted td, which expresses the degree of tax discounting by agents and takes by assumption values between zero and unity. The tax-discounting parameter is defined as the ratio of the future discounted taxes expected by consumers to the value of the tax burden implied by government's current and future discounted liabilities --i.e. the sum of end-of-period government debt and the present value of future government non-interest expenditures. Given the planning horizons of consumers and of the government, the value of td depends on the distribution of the tax burden over time, tending to unity (zero) as more (less) of the tax burden is shifted towards the current generations of consumers. On the other hand, given the distribution of taxes over time, the value of td depends on the degree of coincidence between the planning horizons of consumers and of the government, tending to unity (zero) as consumers' foresight increases (decreases) relative to the government's. Therefore, td cannot be presumed to be constant over time, since the intertemporal

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[10]We call this measure Hicksian because it is consistent with Hicks "central concept of income" as "the maximum value (the agent) could have spent on consumption while maintaining the real amount of his capital stock intact".

distribution of taxes or the foresight of agents may vary, responding to changes in government budgetary policies and to new information available to consumers[11].

Using the parameters hc and td, the wealth and income variables in (1) can be redefined. Indicating by W' real per capita private wealth net of the stock of public debt, by B real per capita government debt, by π the rate of inflation, by S the real per capita government surpluses, by $Y^{d\prime}$ and Y' the conventional measures of real per capita disposable and national income, by T and G real per capita taxes and government consumption and defining the tax-to-income and expenditure-to-income ratios $\theta = T/Y'$, $\phi = G/Y'$, the following expressions can be derived:

$$W_t = W'_t + (1 - td)\ B_t$$

$$Y^d_t = Y^{d\prime}_t - hc\ (1 - td)\ \pi_t\ B_{t-1} + td\ S_t$$

$$\hat{y}_t = \hat{y}'_t + td\ \phi_{t-1}\ (\hat{y}'_t - \hat{g}_t) + (1 - td)\ \theta_{t-1}\ (\hat{y}'_t - \hat{t}_t)$$

Substitution of the above definitions in (1) ensures that, as td varies between zero and unity, the definitions of income and wealth change, implying traditional or Ricardian effects of government debt and deficits on private consumption. When td approaches zero, government debt affects private consumption in the same way as other assets do, while government deficits have the usual positive effects on private consumption through the net government transfer component of disposable income. On the other hand, when td approaches unity, government debt has no influence on private consumption and the appropriate definition of disposable income changes to national income net of government consumption. In this case, the Hicksian-correction factor, hc, is

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[11]See Annex A for details on the definition of td and a derivation of the income and wealth variables defined below.

irrelevant since government transfers are altogether irrelevant for the determination of consumption. Income growth depends on td in a similar way: it is computed net of the growth in taxes when td is zero and net of the growth in government consumption when td is unity.

II.2 Cointegration and error-correction representation

Equation (1) relates equilibrium *per capita* consumption to the contemporaneous levels of several macroeconomic variables. In a stochastic setting, it is natural to expect that the deviations of consumption from its equilibrium value should form a stationary (mean-reverting) series, otherwise the equilibrium specification (1) would be incorrect. Stationarity of the equilibrium errors generally requires that some or all the variables in (1) are cointegrated --since consumption, income, wealth and interest rates are often characterised by unit-root non-stationarity over time[12]. Cointegrated systems allow individual time-series to be integrated, but require certain linear combinations of the series to be stationary or integrated of a lower order[13]. Under the maintained hypothesis that the variables in (1) form a cointegrating system, the equilibrium consumption model can be expressed as a system of simultaneous equations. Defining the matrix of regressors $x_t^{*\prime} = (R_t, w_{t-1}, y_t^d, \hat{y}_t)$, the corresponding (normalised) cointegrating vector $a^{*\prime} = (a_1, a_2, a_3, a_4)$ and the disturbances u_{ct}, $u'_{xt} = (u_{1t}, u_{2t}, u_{3t}, u_{4t})$, we can rewrite (1) as follows:

(1a) $$c_t = a_0 + x_t^{*\prime} a^* + u_{ct}$$

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[12] A growing body of empirical literature has found evidence of unit-root non-stationarity in many macroeconomic time-series [see, for instance, Nelson and Plosser (1982); and Perron (1988)]. Mankiw and Shapiro (1985), Campbell and Deaton (1987) and Stock and West (1987) focus on consumption and income.

[13] If a series x needs to be differenced d times to become stationary, it is called integrated of order d, denoted I(d); a stationary series is I(0). Formal definitions of integration and cointegration in general settings are provided by Granger (1986) and Engle and Granger (1987).

(1b) $\hat{x}^*_t = u_{xt}$

where (u_{ct}, u_{xt}) is generated by a stationary vector ARMA process.

Cointegrating systems such as (1a)-(1b) have several implications for specification, estimation and testing. First, due to the unit-root structure of the regressors, OLS estimates of the cointegrating vector a^* converge to their true values at a much faster rate than ordinary OLS [Stock (1987)]. Second, mainly due to the systemic nature of cointegrating relationships, normal inference cannot be applied to these coefficient estimates, since their asymptotic distributions are non-standard [Phillips and Durlauf (1986)]. However, prior information about non-stationarity and cointegration can be used to design single-equation estimation techniques that improve statistical efficiency, remove bias and bring the asymptotic distributions of the estimated long-run coefficients within the realm of conventional theory, so that usual testing procedures can be applied. A particularly useful technique is the "fully modified least squares" (FMLS) methodology [Phillips and Hansen (1989)], which will be used in the next section.

Cointegrated systems also have a natural error-correction representation, which makes it possible to model explicitly the implicit short-run dynamics contained in the error term u_{ct}. Recently, Engle and Granger (1987) established that if a set of variables is cointegrated there always exists a generating mechanism having the error-correcting form and, conversely, data that is generated by an error-correction model must be cointegrated. In the context of (1a)-(1b) the error-correction mechanism expresses the past deviation of consumption from its long-run (stochastic) equilibrium path and agents adjust consumption in the short-run in order to remove a fraction δ of the resulting disequilibrium each period[14]. The error-correction representation of the long-

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[14]Error-correction models have been successfully used in consumption analysis since the influential work of Davidson et al. (1978). Salmon (1982) and Nickell (1985) show that, under certain conditions, error-correcting behavior

run consumption model (1a)-(1b) can be derived by taking first differences of the cointegrating equation (1a) and substituting from (1b):

$$\hat{c}_t = \delta(x^{*\prime}_{t-1} a^* - c_{t-1}) + v_t$$

with $v_t = u_{ct} + u'_{xt} a^*$.

This error-correction representation can be further simplified imposing the unit-elasticity restriction $a_2 + a_3 = 1$, in line with standard error-correction analyses of consumption. Defining the modified regressor matrix $x'_t = [R_t, (w_{t-1} - c_t), (y^d_t - w_{t-1}), \hat{y}_t]$ and the cointegrating vector $\beta' = \delta a' = \delta(a_1, 1, a_3, a_4)$, we get:

$$\hat{c}_t = x'_{t-1}\beta + v_t$$

Finally, we model the short-run dynamics contained in the error-term v_t as an autoregressive distributed-lag function of consumption and its determinants, obtaining the following general error-correction model[15]:

$$(2) \qquad \hat{c}_t = x'_{t-1}\beta + \Sigma^p_{m=1} \alpha_m \hat{c}_{t-m} + \Sigma^p_{m=0} \theta_m \hat{x}^*_{t-m} + e_t$$

where α_m and the parameter vector θ_m are impact elasticities and the lag-length p and the variables in $\hat{x}^*_{t-m}$ are chosen parsimoniously in order to obtain an i.i.d. disturbance e_t.

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represents the optimal response of economic agents in a dynamic environment.

[15] This "general-to-simple" methodology is exposed, for instance, in Hendry and Richard (1982).

III. Empirical analysis

III.1. The data

We estimate the consumption models (1a)-(1b) and (2) using annual time-series data for Italy and Belgium in the 1953-87 period. The model is estimated using two private consumption aggregates: non durable consumption and total consumption expenditure (as measured by the National Accounts). Disposable income and wealth refer to the private sector. Wealth is obtained as the sum of the stock of capital of the business sector, the housing stock, government debt and net foreign assets. Government debt is defined as outstanding liabilities less financial assets, and relates to the general government for Belgium and to the public sector for Italy. Government surpluses are the net savings of general government. The real interest rate is computed as the difference between a representative medium term nominal yield and the log change of the private consumption deflator. All level variables are in per capita terms and deflated by the private consumption deflator. A description of data sources and definitions is contained in Annex B[16].

Italy and Belgium alternated phases of relatively stable budgetary policies to phases of strong fiscal imbalances over the sample period. Hence, these countries present ideal historical experiments to check if a relationship exists between the degree of tax discounting and the size and dynamics of public debt. Chart 1 presents three indicators of the stance and long-run feasibility of government debt financing policies: the net public debt and public saving ratios to national income, and the spread between the average nominal interest rate on public debt and the rate of growth of nominal

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[16]The data is updated and extended from Nicoletti (1988a). The reader is referred to that study for a discussion of the economic and measurement issues raised by the choice of the above aggregates in the context of studies of tax-discounting behavior.

income[17]. Both countries reached the end of the 1980s with huge levels of public indebtedness, large government dissavings and non-positive spreads between interest and growth. In Italy, chronic government dissavings and an upward trend of the public debt ratio appeared as early as 1964. These trends persisted in spite of a favourable spread between the growth of income and the average interest rate on public debt, suggesting that government dissavings originated mostly from large primary deficits. In Belgium, government budgetary policies generated sizeable surpluses and a declining debt-ratio for most of the sample period; only after 1974 is there an inversion of tendency and public debt starts increasing rather abruptly by 1978. This date is also a watershed for the spread between interest and growth, generally positive before 1978 and strongly negative thereafter. In synthesis, Belgium presents a clear situation in which a sudden deviation from sustainable government financing policies has occurred at the end of the seventies, while in Italy this change in trends has occurred at a much earlier stage of our sample.

III.2. Non-stationarity, cointegration and long-run model estimates

Consistent estimation of the long-run consumption model requires testing for the order of integration and for the cointegration properties of the time-series processes c_t and x_t^*. Unit-root tests and cointegration tests --which check the stationarity of the regression residuals-- can be interpreted as specification tests of model (1a)-(1b).

Unit-root tests are performed using Phillips and Perron (1988) z-statistics and follow the testing strategy proposed by Perron (1988). The z-statistics are transformations of the first-order Dickey and Fuller statistics based --for any time series x-- on the following three alternative regression equations [Fuller (1976), Dickey and Fuller (1981)]:

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[17] The average nominal interest rate on public debt is computed as the ratio of net interest paid by the general government to net public debt outstanding at the beginning of the period.

(i) $\hat{x}_t = \bar{\mu} + \bar{\beta}(t - T/2) + \bar{\alpha}\, x_{t-1} + \bar{u}_t$

(ii) $\hat{x}_t = \mu^* + \alpha^*\, x_{t-1} + u^*_t$

(iii) $\hat{x}_t = \alpha'\, x_{t-1} + u'_t$

The statistics tabulated by Dickey and Fuller are: the point estimates of the autoregressive coefficients $\bar{\alpha}$, α^* and α' of regressions (i)-(iii); the t-statistics, $t_{\bar{\alpha}}$, $t_{\alpha*}$ and $t_{\alpha'}$, associated with these coefficients; and the F-statistics corresponding to the null hypotheses $(\mu^*,\alpha^*) = (0,0)$, $(\bar{\mu},\bar{\beta},\bar{\alpha}) = (0,0,0)$ and $(\bar{\mu},\bar{\beta},\bar{\alpha}) = (\mu,0,0)$ --denoted ϕ_1, ϕ_2 and ϕ_3 respectively. The z-statistics operate non-parametric asymptotic corrections for the nuisance parameters implied by the possible ARMA structure of the disturbances in (i)-(iii)[16]. The testing strategy leads to the following "decision tree":

- Estimate model (i), which includes both a drift and a deterministic trend. If any of the statistics $z(\bar{\alpha})$, $z(t_{\bar{\alpha}})$ and $z(\phi_3)$ is significant, the null of integration is rejected. If not,
- use the test $z(\phi_2)$ [on model (i)] to verify that the drift μ is zero. If this test is significant, it is not possible to switch to model (ii) since the statistics from this model are not invariant to the presence of a drift under the null. Therefore, the procedure ends by acceptance of integration. If ϕ_2 is not significant (i.e. the absence of a drift cannot be rejected),

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[16]The z-statistics present several advantages over the augmented Dickey and Fuller tests: (a) being based on the first-order Dickey and Fuller statistics, they provide an easy and data-saving way to test for unit roots in the case of heterogeneously and dependently distributed data; (b) they can be used with existing tabulations since their asymptotic distributions coincide with those of the simple Dickey and Fuller statistics; (c) they make possible to use the α-statistics even in models with quite general error-structure; and (d) they can be shown to have a power equivalent or superior to that of the augmented Dickey and Fuller tests in regressions with ARMA errors, provided that the MA terms are non-negative

acceptance of the null could be due to low power of the tests from model (i). Therefore,

- estimate model (ii). If any of the statistics $z(\alpha^*)$, $z(t_{\alpha^*})$ and $z(\phi_1)$ are significant, the null of integration is rejected. If not,
- estimate model (iii). If any of the statistics $z(\alpha')$ and $z(t_{\alpha'})$ are significant, integration is rejected. If not, the series is considered non-stationary.

This procedure ensures that the null including a non-zero drift is checked first --this being the situation most likely to occur with many macroeconomic time-series having a secular growth component. However, in some cases it might be sufficient on _a priori_ grounds to begin with model (ii) --as in the case of series already expressed in first differences.

Unit root tests on the variables of model (1a)-(1b) present the non-standard feature that three regressors --w, y^d and $\hat{y}$-- are (non-linear) functions of other variables and of the two unknown parameters hc and td. This entails two complications. First, in the event that the variables composing W, Y^d and $\hat{y}$ were non stationary, some combination of values of the Hicksian-correction and tax-discounting parameters, say (hc*,td*), could make them stationary --i.e. they could be cointegrated with cointegrating vector (hc*,td*). Second, the possible non-stationarity of W and Y^d in absolute real per capita levels does not necessarily imply their non-stationarity in logarithms, since non stationarity is not always preserved by non-linear transformations[19]. Therefore, we followed a testing procedure based on grid-search methods. We first check the order of integration of the real per capita logs of the consumption aggregates c^n and c, the real interest rate R, and of the components of the income growth variable $\hat{y}$. Next, we perform unit-root tests on wealth and disposable income w and y^d, over a grid of six values of hc and td between zero and unity.

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[19] If x is a non-stationary series and f is a non-linear function of x, the series y = f(x) is non stationary only if the function f is asymptotically linear --i.e. if $\lim_{x\to\infty}[f(x)/x] = a$ with $0 < a < \infty$ [see Escribano (1987) and Granger (1986)]. Of course the logarithmic transformation is not asymptotically linear.

The results of the tests are presented in Table 1. Table 2 replicates the tests on the first differences of the series, in order to check for their order of integration. For the series w and y^d, the tables show the minimum and maximum values reached by each of the z-statistics over the grid. The null of non-stationarity cannot be rejected at 5% levels of significance for c^n, c, R, w and y^d. In particular, for w an y^d, the grid-search yields maximum values that are well below the 5% critical values for $z(\tilde{a})$, $z(t_{\tilde{a}})$ and $z(\phi_3)$, and minimum values that are well above the 5% critical value for $z(\phi_2)$[20]. Since $z(\phi_2)$ is always strongly significant, the series are characterized by the presence of growth components, except the real interest rate which is adequately represented by models (ii) and (iii). On the other hand, non-stationarity can easily be rejected for the income growth series --$\hat{y}'$, $\phi(\hat{y}'-\hat{g})$ and $\theta(\hat{y}'-\hat{t})$-- on the basis of model (ii). The tests on the first differences of the non-stationary series are reported in Table 2 [based on model (ii)], and clearly reject non stationarity at very high significance levels. We conclude that the variables in the system (1a)-(1b) are indeed I(1) series, except for the income growth term $\hat{y}$, which is I(0)[21].

Table 3 presents cointegration tests as well as OLS and FMLS estimates of equation (1a). Cointegration tests use two sets of statistics: the cointegrating regression Durbin-Watson (CRDW) and the residual-based versions

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[20]Unit-root tests were also performed on the absolute per capita levels of the series entering the definitions of w and y^d, namely W', B, Y^{d}', πB and S. The test results supported first order non-stationarity of all these series except government debt, which appeared to be I(2). This is consistent with the non-stationarity of the government surplus S, which is an important component of the change in debt. Given the different orders of integration of B and W', total wealth W is necessarily non-stationary. The results of cointegrating tests involving all combinations of the three components of disposable income -- Y^d', S and πB-- excluded the possibility that stationarity of Y^d could be obtained by cointegration of its components.

[21]Earlier studies of the tax-discounting hypothesis in Italy [Modigliani et al. (1985); Modigliani and Jappelli (1987); and Rossi (1988)] ignored the non-stationarity of consumption, income, government deficits and wealth. This cast doubts on the inference procedures followed by these authors.

of the z-statistics for testing unit roots[22]. The FMLS procedure yields estimates that are asymptotically unbiased, efficient and normally distributed by accounting for the endogeneity and serial correlation problems implicit in the system (1a)-(1b) [Phillips and Hansen (1989)]. The model relates I(1) processes to an I(0) process --the income growth variable $\hat{y}$. Gourieroux <u>et al.</u> (1987) show that the properties of OLS estimates of cointegrating vectors remain unaffected by the presence of stationary variables among the regressors. However, the distributions of the z-statistics and of FMLS estimates could be affected. Accordingly, Table 3 presents for each country and for each consumption aggregate two regression estimates: equation 3.1 includes the income growth variable $\hat{y}$, equation 3.2 excludes this variable. Since equation (1a) is non-linear in the two parameters of interest hc and td, we estimate the model by grid-search methods. The outcome of cointegration tests and the model estimates should be considered as conditional on the values of the coefficients hc and td[23].

The outcomes of the cointegration tests are reported at the bottom of the table. The CRDW statistic is generally high, although --with the possible exception of equation 3.1 for the aggregate c-- its value falls in the inconclusive region. The $z(\alpha')$ statistic is always insignificant. However, the $z(t_{\alpha'})$ test rejects the null of no cointegration at 10% levels of significance in most regressions (except equations 3.2 for c^n in Italy, where a 15% significance level is needed for rejection). In many cases this rejection holds at 5% levels of significance (equations 3.1 for c^n in Italy and equations 3.1 for c^n and 3.1, 3.2 for c in Belgium) and in some cases even at 1% levels (equations 3.1, 3.2 for c in Italy). Given the small size of our sample and the generally low power of cointegration tests, these results provide considerable

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[22]CRDW is a finite-sample test proposed and tabulated by Engle and Granger (1987) and Sargan and Bhargava (1983) for the bivariate and multivariate cases respectively; $z(\alpha')$ and $z(t_{\alpha'})$ are asymptotic tests based on the residuals of the cointegrating regression and tabulated by Phillips and Ouliaris (1990).

[23]Non-linear cointegration has received little attention to date [see, however, Escribano (1987)]. The sensitivity of cointegration tests to the values taken by hc and td in the grid search was checked by performing tests for selected values of the couple (hc,td) in the unit simplex (see Table A1 in the Annex).

evidence in favor of cointegration among the I(1) variables of model (1a), especially when the dependent variable is total consumption expenditure c.

The coefficient estimates in Table 3 also tend to support the long-run consumption specification (1a). Coefficients have the expected signs and plausible magnitudes. Income and wealth propensities are in the .55-.85 and .02-.07 ranges, respectively; and the sum of the income and wealth elasticities ranges from .86 to .91[24]. The coefficient of the income growth variable is significant and negative suggesting that dynamic effects may be important in the long-run. Comparison of OLS and bias-corrected FMLS estimates highlights substantial OLS biases in the regressions for c^n in Italy and for c in Belgium, while the bias is negligible in the other cases[25]. FMLS estimates of the wealth and income coefficients are relatively stable across equations, while estimates of the interest rate coefficient are sensitive to the presence of the income growth term. Finally, the grid-search estimates of the Hicksian-correction and tax-discounting parameters hc and td are negligible in most regressions, with the exception of regression 3.1 for c^n in Italy where a sizeable Hicksian-correction effect is detected[26]. The long-run consumption function can therefore be characterized in both countries as a standard "myopic" life-cycle model in which inflation and fiscal illusion prevail and in which government deficits and government debt have the usual positive income and wealth effects. These results are consistent, for Italy, with previous estimates of Hicksian-correction and tax-discounting factors, obtained from static models [see Modigliani, Jappelli and Pagano (1986) and Modigliani and Jappelli (1988)].

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[24] Here and in later sections long-run propensities are computed using the sample averages of c^n, c, y^d and w [with (hc,td) = (0,0)].

[25] Experimental evidence of the finite-sample bias of OLS estimators of cointegrating vectors is provided by Banerjee _et al_. (1986). The superiority of FMLS estimates with respect to simple OLS in finite samples is supported by simulations in Hansen and Phillips (1988).

[26] As shown in Table A1, for Belgium, rejection of the null of no cointegration was obtained only at the values of hc and td resulting from the grid search. For Italy, rejection could be obtained also for (hc,td) = (.5,0) in the c equation and (hc,td) = (1,0) in the c^n equation, suggesting the possibility of a significant Hicksian correction effect in the long-run.

III.3 Error-correction estimates

Table 4 presents maximum likelihood estimates of model (2). The estimates are the result of a simplification-search procedure starting from a "general" formulation of the model, with second order autoregressive and distributed-lag components. Regression 4.1 provides unrestricted estimates of the error-correction component and of the Hicksian and tax-discounting parameters. Regression 4.2 is obtained through the two-step procedure of Engle and Granger (1987), which restricts the error-correction term to be the deviation from the OLS estimates of the long-run model (1a). Finally, regression 4.3 restricts the Hicksian correction parameter to a predetermined value in order to concentrate on the tax-discounting issue. The table also shows diagnostic checks of the various model specifications. These include likelihood-ratio tests of the successive simplifications of the model (LR), Lagrange-multiplier tests of residual autocorrelation (F_{LM}), autoregressive conditional heteroskedasticity tests (F_A), the Jarque and Bera test for normality of residuals (X^2_{JB}) and Chow-tests of structural stability of the variances (F_V) and coefficients (F_C) of the model[27]. In estimating the Hicksian-correction and tax-discounting parameters, hc and td were restricted to lie between zero and unity through a suitable reparameterization of the model, whereby $hc = \sin^2(a_{hc})$ and $td = \sin^2(a_{td})$ [see Box (1966) and Powell (1972)].

In both Italy and Belgium, the error-correction specifications 4.1-4.3 are supported by the data. The error-correction terms are strongly significant and provide additional evidence in favour of the long-run relationship (1a). This is confirmed by the two-step regressions (equation 4.2), in which the lagged OLS residual from the cointegrating regression 3.1 is always strongly

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[27]Details on these tests and on the other statistics shown in the table can be found in Harvey (1981).

significant and correctly signed[26]. The equations for Belgium are particularly satisfactory. The estimated elasticities imply long-run propensities to consume out of income of 0.74 and 0.85 for c^n and c and a wealth propensity of 0.03. Consumers remove around 30 and 40 percent of the previous period disequilibrium in c^n and c, with mean adjustment lags to income of 3.9 and 2.4 years, respectively. On the contrary, equations for Italy yield negative long-run wealth propensities and, especially in the case of c^n, implausibly long mean adjustment lags. However, no sign of mispecification can be detected from the diagnostic tests.

Estimates of the Hicksian-correction and tax-discounting parameters from equation 4.1 are strongly significant and quite close to unity in both countries. These results are consistent with earlier estimates of hc and td obtained by Nicoletti (1988b), Onofri (1987) and Rossi (1988) using simpler dynamic models of consumption, but contrast strikingly with the static estimates obtained in the previous subsection. This contrast suggests that the Hicksian-correction and tax-discounting effects are closely linked to the short-run adjustment components of consumption behavior. Such an interpretation is corroborated by the results of the two-step estimates: equation 4.2 yields estimates of both hc and td that are much higher than the corresponding estimates obtained from the cointegrating regressions 3.1-3.2, although hc and td are in fact restricted to zero in the error-correction term. For Italy, the two-step estimates of td are 0.49 for c^n and 0.61 for c, while for Belgium td rises from zero to 0.26 and 0.29 for c^n and c respectively (for Belgium, no estimates of hc could be obtained from equation 4.2 since this parameter is not contained in the short-run components of the equation). In all cases, the estimates of hc and td are strongly significant. The restriction of the Hicksian-correction parameter to unity in equation 4.3 cannot be rejected by the likelihood-ratio test and is therefore maintained for the rest of the analysis.

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[26] If the variables in model (1a) were not cointegrated, regressions 4.1-4.3 would relate processes having different orders of integration and the t-statistics associated with lagged levels of the regressors would tend to be insignificant.

The divergence between FMLS and error-correction estimates indicates that the explicit consideration of short-run dynamics is an important source of identification for tax-discounting effects on consumption. These effects seem to be generated mainly in the process of adjustment of consumption to its long-run equilibrium level. A possible explanation of this phenomenon is that the positive correlation between private saving and government deficits arises only as a reaction to government financing policies that are expected to impose extraordinary burdens on consumers in the short-term because they are perceived as deviations from "normality". If this interpretation was correct, we would expect the long-run consumption specification (1a) to be rather stable over time, while the error-correction specification (2) could be affected by "transitory" movements in public debt dynamics.

Chart 2 graphically investigates the stability of the estimated cointegrating vectors by plotting the one-step residuals obtained from recursive (grid-search) estimation of regressions 3.1 (Table 3). The lack of evident predictive failure in the pattern of the residuals suggests that the estimates are not significantly affected by the choice of the sample period, lending support to our previous characterization of long-run consumption as being life-cycle with no tax-discounting effects. Unfortunately, no formal stability test can be performed on the long-run relationship between consumption, interest rates, income and wealth, since the OLS residuals from cointegrating regressions generally cannot be expected to be i.i.d..

Chart 3 investigates the stability of the short-run consumption model. Quandt's ratio --computed through recursive estimation of the error-correction equations 4.3-- is plotted against several possible breakpoints within the sample[29]. The graphs for Belgium clearly point out that, after a period of

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[29]Quandt's ratio relates the unrestricted sum of squared residuals --obtained running separate regressions before and after the candidate breakpoint-- to the restricted sum of squares obtained over the whole sample [Brown et al. (1975)]. Although this statistic offers no formal test for the identification of the breakpoints, it sheds light on the stability of the regression and the periods in which the ratio attains its lowest levels are plausible candidates for a

remarkable stability, the equations suffer from an abrupt structural break in 1978-79. The graphs for Italy deliver a more complex picture: overall the equations are less stable and it is difficult to pinpoint clear structural breaks, although 1971 and 1976-78 seem to be possible candidates. The Chow-tests in Table 4 confirm this graphical evidence. The results of the F_c-tests indicate the presence of a 1978 break in Belgium and of two breaks in 1971 and in 1977 (for c only) in Italy. Inspection of the recursive coefficient estimates of equation 4.3 suggested that the 1977 break for Italy was caused by a change in the wealth and income elasticities. Allowing for different elasticities in the subperiods 1955-76 and 1977-86 considerably improved the fit of the regression, yielding sensible long-run propensities and plausible mean adjustment lags, while leaving the other coefficient estimates unchanged[30]. On the other hand, it was not possible to relate the 1971 break for Italy and the 1978 break for Belgium to changes in the estimated elasticities, which appear to be rather stable during these periods.

To what extent can these breaks be attributed to changes in government financing policies and to their effects on tax discounting? Chart 4 shows the graphs of the recursive estimates of the tax-discounting parameter for Italy and Belgium. The graphs are striking. Both in Italy and in Belgium, tax-discounting effects begin to appear in correspondence to the changes in government financing policies highlighted in Chart 1. In Belgium, td rises from zero to 0.4 in the 1978-80 period and gradually reaches values close to unity by 1987. In Italy td jumps close to unity in 1967 and remains quite stable thereafter. Table 5 presents the results of various attempts at modelling this time-varying behavior of tax-discounting. The analysis is based on equation 4.3. However, for Italy, we allow for different income and wealth elasticities

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structural break. The ratios in Chart 3 were computed from recursive grid-search estimates of equations 4.3 over a grid of 11 values of td in the unit interval.

[30] In the two subperiods the estimated income propensities for Italy range from 0.63 to 0.73, while the wealth propensities range from 0.01 to 0.05. The mean adjustment lags of consumption to income are about 3.5 years for c^n and 2 years for c. The estimates of td remain close to unity for both consumption aggregates.

over the 1955-76 and 1977-86 subperiods, in order to control for this independent source of instability. For each country, we report the full-sample maximum-likelihood estimates of td and the estimates obtained by introducing dummies that isolate the influence of tax-discounting before and after the breaks evidentiated by Chart 4 (the estimates of the other coefficients of the model are omitted since they remain practically unaffected). Accordingly, the time of the break is set at 1967 for Italy and 1978 for Belgium. The likelihood-ratio statistic tests the null of parameter stability across the subperiods.

The results for Italy and Belgium are quite different. In Italy, the estimates of td remain stable around unity in the two subperiods and the null of parameter stability cannot be rejected by the likelihood-ratio test. This result apparently contradicts the 1967 jump in td evidentiated by the recursive coefficient estimates of Chart 4. However, the period of "stable" financing policies covered by the Italian sample is probably too short to make it possible to identify separate tax-discounting effects before and after the change in policies. On the other hand, rather than reflecting a break in aggregate consumption behaviour, the 1967 jump in td could be caused by a normal instability of the coefficient estimates, due to the small number of observations used in the initial recursions. Therefore, these results indicate the need for further research, possibly based on a longer time-span. In Belgium, the estimates differ widely before and after 1978, being two to three times larger in the later period, and parameter stability can be rejected at 1 percent confidence levels.

The results for Belgium support the hypothesis of a relationship between tax-discounting behavior and changes in government financing trends. In order to explore further this possibility, we explicitly model the behavior of td over time, expressing tax discounting as a non-linear function of the ratio of public debt to national income. This ratio is a natural indicator of trends in government financing policies, being at the same time a synthetic measure of past policies and a conditioning variable for the future. The chosen function is $td = 1 - \exp[-a(B/Y')]$, where a is the parameter to be estimated. This transformation preserves the a priori restriction that the values of the tax-discounting parameter be in the unit interval. Table 5 reports, for each consumption aggregate, the full-sample estimates of a, the dummy-variable estimates of a before and after the 1978 break and the likelihood-ratio tests of parameter stability over the two subsamples.

The dependence of tax-discounting effects on the public debt ratio cannot be rejected by the data. The strong significance of the full-sample estimates of a confirms the observed positive correlation between td and (B/Y'). At the same time, the estimates of a are larger and more significant in the 1978-87 period (a2), and the hypothesis of parameter stability is strongly rejected in the case of consumption of non durables. This suggests that the dependency of td on the debt-ratio is a phenomenon arising only when, as in 1978, there is an abrupt inversion of tendency in the time behavior of this indicator. The temporal patterns of td implied by the full-sample and dummy-variable estimates of the non-linear tax-discounting model are illustrated in Chart 5. Both graphs in the chart show relatively low and declining tax-discounting effects before 1978 and increasing effects after that date. The lower graph evidentiates --especially for the aggregate c^n-- the "trigging" influence of the change in debt-ratio dynamics on the tax-discounting parameter.

Summary and conclusions

If consumers discount future taxes, public debt formation is partially, or totally, compensated by increased private sector savings and has no consequences for long-run capital formation. The evidence for Italy and Belgium suggests that any such positive correlation between government deficits and private saving is tightly linked to the short-run adjustment components of consumption behavior. More importantly, the relationship between debt, deficits and private consumption is not necessarily stable over time and may well depend on the characteristics of the financing regime followed by the government. In both countries, tax-discounting is found to vary over time in coincidence with sharp changes in the trends followed by government financing policies. In particular, the results for Belgium show that strong Ricardian effects appear only after government policies have switched from a stable to an apparently unstable course --characterised by chronic government deficits and rapidly rising stocks of debt. Moreover, when the debt dynamics are explosive, tax-discounting can be modelled as an increasing function of the debt-ratio. A possible interpretation of these results is that tax-discounting is a contingent phenomenon, caused by the fear of unstable policies and of the future cost that the return to stability may impose to consumers in terms of increased tax burdens.

ANNEX A

Derivation of expressions for income and wealth embodying Ricardian effects

We use the following short-hand notation (for any variable x):

$$x_t^H = \Sigma_{i=0}^{H-t} \beta_{t+i} x_{t+i} \qquad \text{with } \beta_{t+i} = \Pi_{s=0}^{i} [1/(1+R_{t+s})]$$

$$x_{t+1}^H = x_t^H - x_t$$

The starting points are the definition of consumers' intertemporal wealth (Q),

A1 $$Q_t = W_{t-1}(1 + R_t) + Y_t^{HC} - T_t^{HC},$$

and the definition of the government's intertemporal budget constraint:

A2 $$T_t^{HG} = B_{t-1}(1 + R_t) + G_t^{HG} \ ,$$

where HC and HG (HC ≤ HG) are the planning horizons of consumers and of the government, respectively.

We now define the tax-discounting parameter td as the ratio of the present value of future taxes expected by consumers to the present value of the tax burden implied by the government's current and future liabilities:

A3 $$td = T^{HC}_{t+1}/T^{HG}_{t+1}$$

Given HC and HG, the value of td depends on the distribution of the tax burden over time; given this distribution, td depends on the degree of coincidence between the planning horizons of consumers and of the government.

Using A2 to eliminate T^{HG}_{t+1} from A3 and rearranging, we get the following expression for the tax burden expected by consumers:

A4 $$T^{HC}_{t} = td\ B_{t-1} + T_t - td\ (S_t + \pi_t B_{t-1}) + td\ G^{HG}_{t+1}$$

Finally, substituting A4 in A1 and allowing for inflation illusion in evaluating the real income effects of government's interest transfers, we obtain the following expression for the intertemporal wealth of consumers:

A5 $$Q_t = [W'_{t-1} + (1 - td)B_{t-1}] + [Y^{d'}_t + td\ S_t - hc(1 - td)\pi_t B_{t-1}] + + [Y'^{HC}_{t+1} - T^{HC}_{t+1} + td(T^{HG}_{t+1} - G^{HG}_{t+1})]$$

The first two expressions in brackets are the definitions of non-human wealth and disposable income used in empirical analysis. The third expression in brackets, which represents human wealth, was approximated for estimation purposes by the current rate of growth in $Y_t = Y'_t - T_t + td(T_t - G_t)$:

$$\hat{y}_t = \hat{y}'_t + td\ \phi_{t-1}\ (\hat{y}'_t - \hat{g}_t) + (1 - td)\ \theta_{t-1}\ (\hat{y}'_t - \hat{t}_t)$$

ANNEX B

Data sources and definitions

The basic series used in the analysis are:

C^d = private consumption of durable goods
C = total private consumption
Y' = national income
S = general government net saving
IG = general government net interest payments
G = general government purchases of goods and services
B = public debt
KBV = real business sector capital stock
KHV = real stock of residential buildings
KF = net foreign assets
NR = long-term nominal interest rate
PC = private consumption deflator
PIB = deflator of gross private capital formation (non-residential)
PIH = deflator of gross private capital formation (residential)
POP = population

The following variables were computed:

$C^n = C - C^d$ = private consumption of non durable goods
T = S + G + IG = taxes net of transfers
$Y^{d}{'}$ = Y' - T + IG = private disposable income
KB = KBV*PIB
KH = KHV*PIH
W' = KB + KH + KF = private wealth
R = NR - $((PC/PC_{-1}) - 1)$ = real interest rate

All variables are deflated by PC and expressed per unit of POP.

ITALY

Italian National Accounts were revised in 1985, no new data going back to the 1950s is yet available. Therefore, the data for C^d, C and Y' refers to the old system of National Accounts. The 1986-87 values of these variables are estimates based on the rates of growth of the corresponding series in new National Accounts data. In many cases, the 1952-59 period was covered using the data of Modigliani, Jappelli and Pagano (1985), which was kindly provided by the authors. Original sources are as follows:

C^d, C, Y', G, PC, PIB, PIH: OECD National Accounts.

S, IG: 1952-79 Annuario di Contabilità Nazionale, ISTAT;
1980-87 Banca d'Italia, Relazione Annuale, Appendice Statistica.

NR: IMF, International Financial Statistics (medium term government bond yield)

POP: United Nations Demographic Yearbook (mid-year estimates).

KBV, KHV: "Ricostruzione di serie storiche settoriali dell'economia italiana", A. Heimler and C. Milana, Consiglio Nazionale delle Ricerche, Progetto finalizzato economia, Working Paper, 1986.

KF: 1952-75 "Il sistema degli stati patrimoniali per l'economia italiana (1948-81)", G. Della Torre, Studi e informazioni, Banca Toscana, 1984;
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B: 1950-59 estimated using rates of growth of central government debt;
1960-84 "L'indebitamento pubblico in Italia: evoluzione, prospettive, problemi", Rapporto alla V Commissione della Camera dei Deputati, L. Spaventa, G. Morcaldo e P. Zanchi, 1984;

1985-87 Banca d'Italia, Relazione Annuale, Appendice Statistica.

BELGIUM

KF: 1950-59 estimated cumulating current account balances from OECD National Accounts;

1960-84 "De Financielle Rekeningen en Stroomtabellen van Belgie 60-84", Planning Papers, Bureau du Plan, Bruxelles;

1985-86 estimated cumulating current account balances from OECD National Accounts.

POP: United Nations Demographic Yearbook (mid-year estimates).

The other series were kindly provided on diskette by Bureau du Plan, Bruxelles.

TABLE A1
Sensitivity of cointegration tests to (hc,td)

Country		ITALY			BELGIUM		
Statistic		CRDW	$z(\hat{\alpha})$	$z(t\hat{\alpha})$	CRDW	$z(\hat{\alpha})$	$z(t\hat{\alpha})$
Dep.var.	(hc,td)						
c^n	(0,0)	.98	-20.94	-3.64	1.57	-20.73	-4.97
	(0,.5)	.87	-17.86	-3.35	.98	-16.94	-3.59
	(0,1)	.76	-14.86	-2.91	.91	-15.78	-3.21
	(.5,0)	1.37	-25.38	-4.46	1.36	-18.73	-3.9
	(.5,.5)	1.01	-19.5	-3.61	.94	-16.27	-3.5
	(1,0)	1.62	-26.02	-4.82	1.16	-16.7	-3.24
c	(0,0)	1.71	-29.83	-5.39	1.71	-24.91	-4.85
	(0,.5)	1.44	-25.74	-4.57	1.16	-19.65	-3.99
	(0,1)	1.12	-19.21	-3.55	.88	-15.31	-3.14
	(.5,0)	1.95	-19.29	-5.96	1.51	-22.4	-4.09
	(.5,.5)	1.56	-25.84	-4.75	1.12	-18.68	-3.83
	(1,0)	1.89	-27.18	-5.5	1.27	-18.89	-3.34

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TABLE 1

Unit-root tests: ITALY (1954-1986)

Number of observations: 33

Order of truncation lag: 4 (a)

Variables (c)		Test-statistics (b) $\hat{\alpha}$	$z(\hat{\alpha})$	$z(t_{\hat{\alpha}})$	$z(\Phi_3)$	$z(\Phi_2)$	α^*	$z(\alpha^*)$	$z(t_{\alpha^*})$	$z(\Phi_1)$	α'	$z(\alpha')$	$z(t_{\alpha'})$
c^n		1.01	-0.56	-0.3	1.81	23.35							
c		1.01	-0.43	-0.23	1.58	22.46							
R		0.83	-6.69	-1.6	1.23	1.18	0.82	-7.26	-1.87	1.88	-0.17	-6.78	-1.81
$\dot{y}'$							0.35	-23.4	-4.05	8.73			
$\phi(\dot{y}'-\dot{g})$							-0.05	-32.08	-6.17	16.93			
$\theta(\dot{y}'-\dot{\epsilon})$							1	-36.67	-5.58	16.98			
w	max		-7.79	-1.95	1.64	17.75							
	min		-5.14	-1.33	0.85	9.83							
y^d	max		-1.86	-0.94	2.6	19.44							
	min		-0.19	-0.11	1.76	12.63							

Notes: (a) The results remain unchanged when the lag is set to 8.
(b) The z-statistics have the same critical values as the simple Dickey and Fuller statistics [Fuller (1976), pp.371,373; Dickey and Fuller (1981), pp.1062-1063].
(d) c, c^n, w, y^d in real per capita logs; $\dot{y}'$, $\dot{g}$, $\dot{\epsilon}$ are log changes of real per capita levels; R is in percent.

TABLE 1 (continued)

Unit-root tests: BELGIUM (1956-1987)

Number of observations: 32

Order of truncation lag: 4 (a)

Variables (c)		Test-statistics (b)											
		α	$z(\alpha)$	$z(t_\alpha)$	$z(\Phi_3)$	$z(\Phi_2)$	α^*	$z(\alpha^*)$	$z(t_{\alpha^*})$	$z(\Phi_1)$	α'	$z(\alpha')$	$z(t_{\alpha'})$
c^d		0.91	-6.27	-1.72	1.48	17.05							
c		0.94	-4.9	-1.41	1.04	14.29							
R		0.56	-15.4	-3.05	3.45	3.45	0.6	-14.68	-2.9	4.49	-0.16	-4.63	-1.55
$\hat{y}'$							0.23	-29	-4.49	7.89			
$\phi(\hat{y}'-\hat{g})$							0.17	-27.8	-4.77	7.8			
$\theta(\hat{y}'-\hat{t})$							-0.4	-39.3	-9.45	12.26			
w	max		-10.9	-2.38	2.04	36.93							
	min		-2.37	-0.82	1.29	11.52							
y^d	max		-10.07	-2.31	2.01	8.01							
	min		-4.27	-1.34	0.95	5.17							

Notes: (a) The results remain unchanged when the lag is set to 8.
(b) The z-statistics have the same critical values as the simple Dickey and Fuller statistics [Fuller (1976), pp.371,373; Dickey and Fuller (1981), pp.1062-1063].

(d) c, c^d, w, y^d in real per capita logs; $\hat{y}'$, $\hat{g}$, $\hat{t}$ are log changes of real per capita levels; R is in percent.

TABLE 2

Unit-root tests on first differences
Order of truncation lag: 4 (a)

Sample		ITALY				BELGIUM			
Test-statistics (b)		α^*	$z(\alpha^*)$	$z(t_{\alpha^*})$	$z(\phi_1)$	α^*	$z(\alpha^*)$	$z(t_{\alpha^*})$	$z(\phi_1)$
Variables (c)									
c^*		0.55	-16.12	-3.16	5.36	0.34	-21.62	-3.87	7.76
c		0.48	-17.93	-3.4	6.13	0.34	-21.67	-3.85	7.7
R		0.14	-24.46	-4.69	9.36	-0.34	-41.26	-7.85	30.44
w	max		-22	-4.5	7.85		-24.75	-4.89	9.77
	min		-19.85	-4.24	6.08		-16.68	-3.29	5.39
y^d	max		-27.94	-4.36	10.49		-46.05	-7.45	31.01
	min		-23.39	-3.96	8.36		-30.72	-4.65	12.73

Notes: See Table 1.

TABLE 3
Cointegrating regressions (a): ITALY (1953-1986)

Dep. var.	c^*				c			
Regression (b)	3.1		3.2		3.1		3.2	
Procedure	OLS	FMLS	OLS	FMLS	OLS	FMLS	OLS	FMLS
Regressor (c)								
const.	-1.56	-1.58	-1.65	-1.64	-1.35	-1.36	-1.43	-1.44
		(-17.35)		(-13.19)		(-21.26)		(-16.8)
R	0.05	0.15	0.13	0.04	0.05	0.03	0.02	0.005
		(1.93)		(0.33)		(0.47)		(0.06)
w	0.13	0.21	0.19	0.15	0.13	0.14	0.18	0.19
		(6.53)		(3.27)		(5.19)		(5.77)
y^d	0.71	0.64	0.66	0.69	0.73	0.72	0.68	0.67
		(17.72)		(13.5)		(23.99)		(18.16)
$\hat{y}$	-0.35	-0.26			-0.29	-0.25		
		(-2.84)				(-4.02)		
hc	0.5	0.8	0.4	0.2	0.1	0.1	0	0
td	0	0	0	0	0	0	0	0
Coint. tests (d):								
CRDW	1.37		1.23		1.79		1.66	
$z(\alpha')$	-25.38		-20.73		-30.02		-26.65	
$z(t_{\alpha})$	-4.46 **		-3.82		-5.61 ***		-4.82 **	

Notes: (a) Estimated by grid-search over 121 values of (hc,td) in the unit simplex.
(b) The truncation lag for FMLS estimates is 4; estimates remain unchanged with an 8th order lag.
(c) T-statistics in parenthesis.
(d) Critical values for CRDW can be found in Sargan and Bhargava (1983); critical values for the z-tests are provided by Phillips and Ouliaris (1990). An asterisk indicates significance at 15%, two asterisks at 5% and three asterisks at 1%.

TABLE 3 (continued)
Cointegrating regressions (a): BELGIUM (1955-1987)

Dep. var.	c^*				c			
Regression (b)	3.1		3.2		3.1		3.2	
Procedure	OLS	FMLS	OLS	FMLS	OLS	FMLS	OLS	FMLS
Regressor (c)								
const.	-0.66	-0.7	-0.83	-0.84	-0.4	-0.44	-0.51	-0.56
		(-17.5)		(-11.59)		(-9.96)		(-9.14)
R	0.33	0.22	0.1	-0.07	0.04	0.005	-0.1	-0.25
		(3.29)		(-0.57)		(0.07)		(-2.4)
w	0.18	0.21	0.31	0.32	0.16	0.2	0.24	0.3
		(7.23)		(6.01)		(6.14)		(6.53)
y^d	0.68	0.64	0.55	0.53	0.74	0.71	0.66	0.61
		(20.59)		(9.35)		(20.4)		(12.41)
$\dot{y}$	-0.4	-0.37			-0.25	-0.3		
		(-9.49)				(-7.05)		
hc	0	0	0	0	0	0	0	0
td	0	0	0	0	0	0.1	0	0.1
Coint. tests (d):								
CRDW	1.57		1.23		1.71		1.57	
$z(\alpha')$	-20.73		-22.09		-24.91		-28.59	
$z(t_{\alpha})$	-4.97 **		-4.04 *		-4.85 **		-4.7 **	

Notes:
(a) Estimated by grid-search over 121 values of (hc,td) in the unit simplex.
(b) The truncation lag for FMLS estimates is 4; estimates remain unchanged with an 8th order lag.
(c) T-statistics in parenthesis.
(d) Critical values for CRDW can be found in Sargan and Bhargava (1983); critical values for the z-tests are provided by Phillips and Ouliaris (1990). An asterisk indicates significance at 15%, two asterisks at 5% and three asterisks at 1%.

TABLE 4
Error-correction regressions:ITALY (1955-86)
Maximum likelihood estimates

Dependent variable:	$\hat{c}^n$			$\hat{c}$		
Equation(a):	4.1	4.2	4.3	4.1	4.2	4.3
Regressor(b):						
const.	0.1	0.02	0.1	0.002	0.02	-0.00
	(2.83)	(5.5)	(3.21)	(0.05)	(6.46)	(-0.01)
r_{t-1}	0.21		0.21	0.16		0.16
	(4.52)		(5.7)	(4.1)		(5.1)
w_{t-2}/c_{t-1}	0.11		0.12	0.27		0.28
	(2.9)		(3.66)	(5.74)		(6.75)
z_{t-1}		-0.28			-0.54	
		(-2.58)			(-4.24)	
y^d_{t-1}/w_{t-2}	0.19		0.2	0.28		0.29
	(3.95)		(6.47)	(6.14)		(9.2)
$\hat{y}_{t-1}$	-0.01			-0.00		
	(-0.34)			(-0.00)		
$\hat{w}_t$	-0.1	-0.13	-0.09	-0.07	-0.07	-0.07
	(-2.51)	(-2.32)	(-2.38)	(-1.87)	(-1.31)	(-1.91)
$\hat{y}_{t-1}$	0.28	0.35	0.28	0.32	0.38	0.32
	(8.17)	(10.08)	(8.3)	(10.99)	(11.48)	(11.57)
$\hat{c}_{t-2}$	-0.29	0.1	-0.3	-0.44	-0.1	-0.45
	(-2.62)	(1.02)	(-3.42)	(-5.25)	(-1.24)	(-6.11)
$\hat{y}^d_{t-2}$	0.2	0.19	0.19	0.27	0.27	0.27
	(4.22)	(3.56)	(4.46)	(6.4)	(5.24)	(6.39)
$hc=\sin^2(a_{hc})$	1	1	1	1	1	1
	(1.58)	(3.73)		(1.41)	(18.45)	
$td=\sin^2(a_{td})$	0.94	0.61	0.94	0.98	0.49	0.98
	(2.87)	(4.26)	(3.03)	(3.24)	(5.0)	(2.92)

TABLE 4
Error-correction regressions:ITALY (1955-86)
Maximum likelihood estimates

Statistics(c):						
Degrees of freedom (N)	20	23	23	20	23	23
S.E.%	1.05	1.2	0.98	0.97	1.14	0.98
LR	5.22		5.32	7.56		7.56
$F_{LM}(4,N-4)$	0.11	1.23	0.11	0.62	1.18	0.64
$F_{LM}(1,N-1)$	0.05	0.23	0.08	0.46	1.09	0.54
$F_A(2,N-2)$	1.51	2.37	1.54	0.9	1.51	0.84
X^2_{JB}	.88	.8	1.28	.39	.6	.55
$F_V(2,14)$			0.26			0.37
$F_C(8,16)$			1.36			2.7
$F_V(9,16)$			1.52			1.7
$F_C(8,16)$			5.54*			6.18*

Notes: t-statistics in parenthesis; for hc and td the t-statistics refer to a_{hc} and a_{td} in $hc=\sin^2(a_{hc})$ and $td=\sin^2(a_{td})$.
(a) hc restricted to unity in regression 6.3
(b) z is the residual from the cointegrating regression 3.1 (Table 3).
(c) LR denotes the $X^2(p)$ likelihood-ratio test associated with p parameter restrictions; S.E. is the adjusted standard error of the regression; $F_{LM}(p,N-p)$ is the modified Lagrange-multiplier test for serial correlation up to the p-th order; $F_A(p,N-p)$ is the modified Lagrange-multiplier p-th order ARCH test; $X^2_{JB}(2)$ is the Jarque and Bera test for normality of residuals; $F_V(n_1-K,n_2-K)$ and $F_C(K,T-2K)$ are Chow-tests for the constancy of the variances and of the K coefficients of the equation over the two subsamples of size n_1 and n_2 ($n_1+n_2=T$). An asterisk indicates significance at 5%.

TABLE 4
Error-correction regressions: BELGIUM (1956-87)
Maximum likelihood estimates

Dependent variable	$\hat{\ell}^n$			$\hat{\ell}$		
Equation(a):	4.1	4.2	4.3	4.1	4.2	4.3
Regressor(b):						
const.	-0.05	0.01	-0.06	-0.03	0.01	-0.03
	(-2.93)	(4.42)	(-3.93)	(-1.29)	(4.06)	(-1.71)
r_{t-1}	0.31		0.23	0.16		0.13
	(5.29)		(3.94)	(2.48)		(2.61)
w_{t-2}/c_{t-1}	0.36		0.33	0.44		0.43
	(11.99)		(9.1)	(10.11)		(10.78)
z_{t-1}		-0.85			-0.96	
		(-6.85)			(-6.81)	
y^d_{t-1}/w_{t-2}	0.34		0.29	0.4		0.38
	(10.48)		(8.21)	(8.63)		(9.6)
$\hat{y}_{t-1}$	-0.09			-0.04		
	(-1.9)			(-0.82)		
$\hat{y}_{t-1}$	0.32	0.32	0.31	0.45	0.45	0.45
	(10.38)	(8.27)	(9.21)	(13.46)	(12.4)	(13.38)
$\hat{\ell}_{t-1}$	-0.36	0.2	-0.4	-0.28	0.25	-0.31
	(-4.12)	(2.49)	(-4.27)	(-3.86)	(3.5)	(-4.81)
$\hat{\ell}_{t-2}$	-0.26	0.004	-0.26	-0.23	-0.03	-0.23
	(-3.66)	(0.05)	(-3.45)	(-3.49)	(-0.51)	(-3.39)
$hc=\sin^2(a_{hc})$	0.95		1	1		1
	(12.52)			(2.62)		
$td=\sin^2(a_{td})$	1	0.26	1	0.73	0.29	0.74
	(3.11)	(2.57)	(4.63)	(7.98)	(5.04)	(7.81)

TABLE 4
Error-correction regressions: BELGIUM (1956-87)
Maximum likelihood estimates

Statistics (c):						
Degrees of freedom (N)	21	25	24	21	25	24
S.E.%	0.97	1.03	0.94	0.88	1	0.83
LR	3.43		6.43	4.03		4.55
$F_{LM}(4,N-4)$	1.89	1.18	0.81	1.52	0.61	1.97
$F_{LM}(1,N-1)$	1.19	2.33	0.07	0.94	0.01	0.24
$F_A(2,N-2)$	0.68	0.75	0.48	0.87	1.13	0.87
X^2_{JB}	1.74	.33	1.18	1.16	.32	1.18
$F_V(2,15)$			0.07			0.06
$F_C(7,17)$			2.72*			2.89*
$F_V(10,7)$			0.79			0.62
$F_C(7,17)$			0.91			1.49

Notes: t-statistics in parenthesis; for hc and td the t-statistics refer to a_{hc} and a_{td} in $hc=\sin^2(a_{hc})$ and $td=\sin^2(a_{td})$.

(a) hc restricted to unity in regression 6.3

(b) z is the residual from the cointegrating regression 3.1 (Table 3).

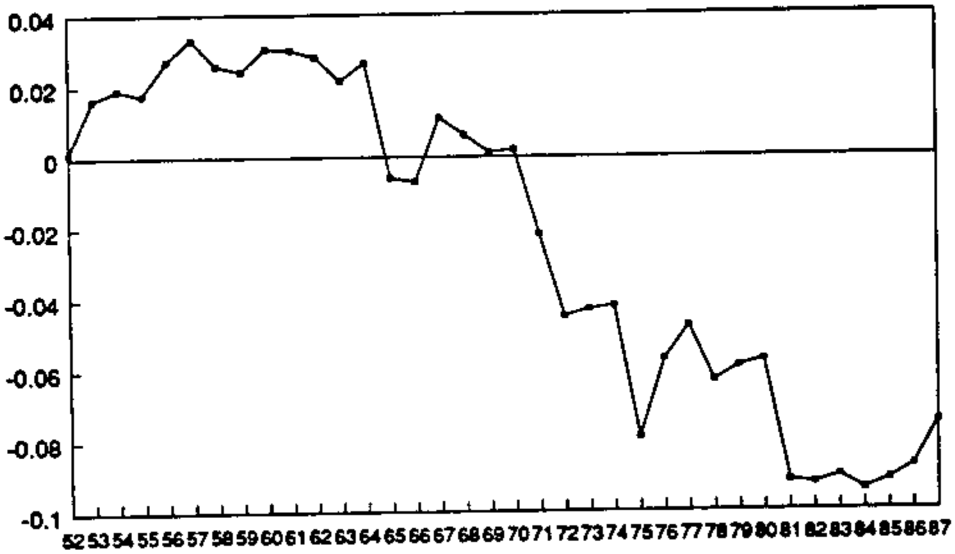
(c) LR denotes the $X^2(p)$ likelihood-ratio test associated with p parameter restrictions; S.E. is the adjusted standard error of the regression; $F_{LM}(p,N-p)$ is the modified Lagrange-multiplier test for serial correlation up to the p-th order; $F_A(p,N-p)$ is the modified Lagrange-multiplier p-th order ARCH test; $X^2_{JB}(2)$ is the Jarque and Bera test for normality of residuals; $F_V(n_1-K,n_2-K)$ and $F_C(K,T-2K)$ are Chow-tests for the constancy of the variances and of the K coefficients of the equation over the two subsamples of size n_1 and n_2 $(n_1+n_2=T)$. An asterisk indicates significance at 5%.

TABLE 5
Time-varying tax-discounting
Maximum-likelihood estimates and stability tests (a)

Sample	Italy (1955-1986)				Belgium (1956-1987)							
Tax-discounting model	td=$\sin^2$(a)				td=$\sin^2$(a)				td=1-exp(-a(b/y))			
Dependent variable	c^*		c		c^*		c		c^*		c	
Time of break	no	1967	no	1967	no	1978	no	1978	no	1978	no	1978
Estimated parameter (a)												
a									1.89		1.29	
									(3.47)		(3.98)	
td	1		0.92		1		0.74					
	(2.36)		(6.36)		(4.63)		(7.81)					
a1										0.31		0.85
										(1.13)		(2.25)
td1		1		1		0.35		0.4				
		(3.32)		(3.21)		(4.27)		(4.18)				
a2										2.49		1.56
										(2.46)		(3.43)
td2		0.94		0.89		1		0.79				
		(2.86)		(6.74)		(3.75)		(9.53)				
LR		0.08		0.44		11.62		7.13		8.16		1.74

Notes: (a) Based on equation 4.3. For Italy, the equation allows for different income and wealth elasticities in the subperiods 1955-76 and 1977-86.
(b) a and td are the full sample estimates; a1, a2 and td1, td2 are the estimates before and after the breakpoints.

CHART 1
Net public saving ratios
ITALY (1952-87)



BELGIUM (1953-87)

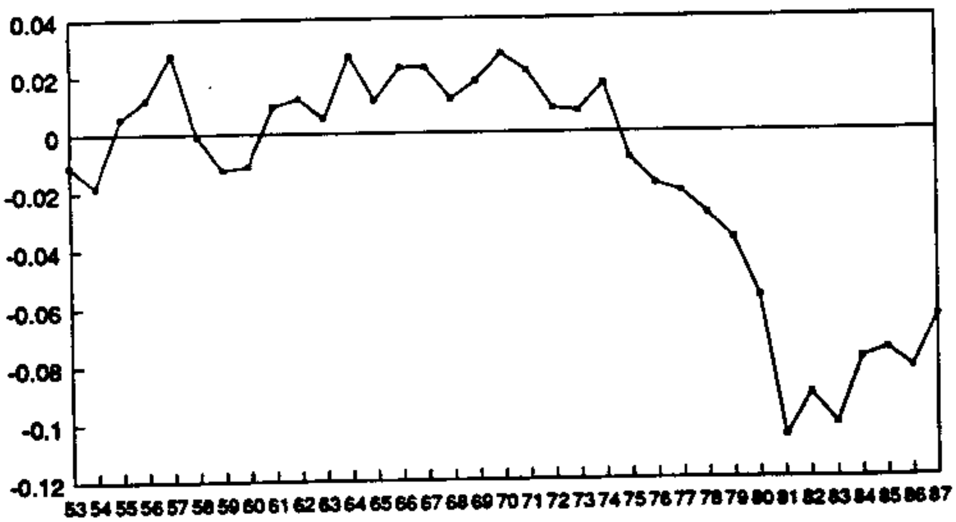
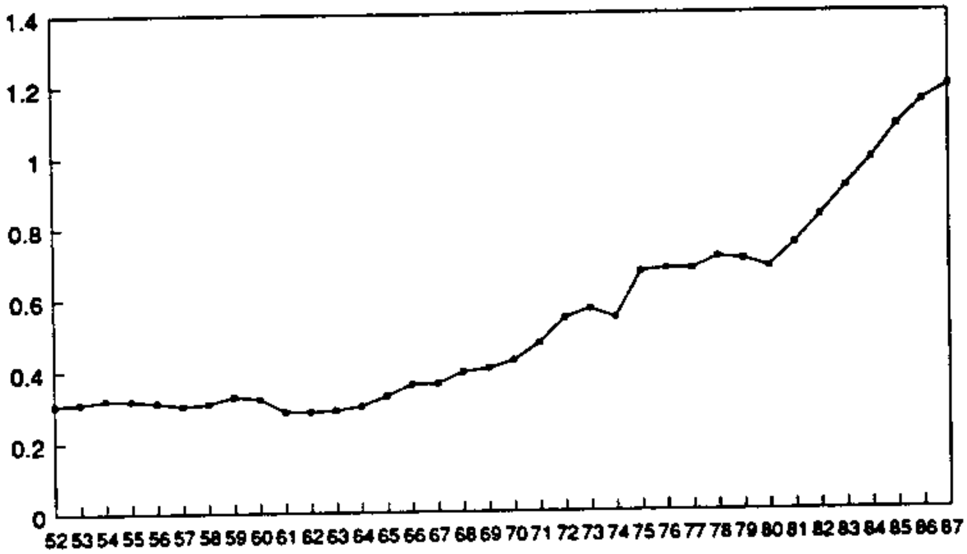
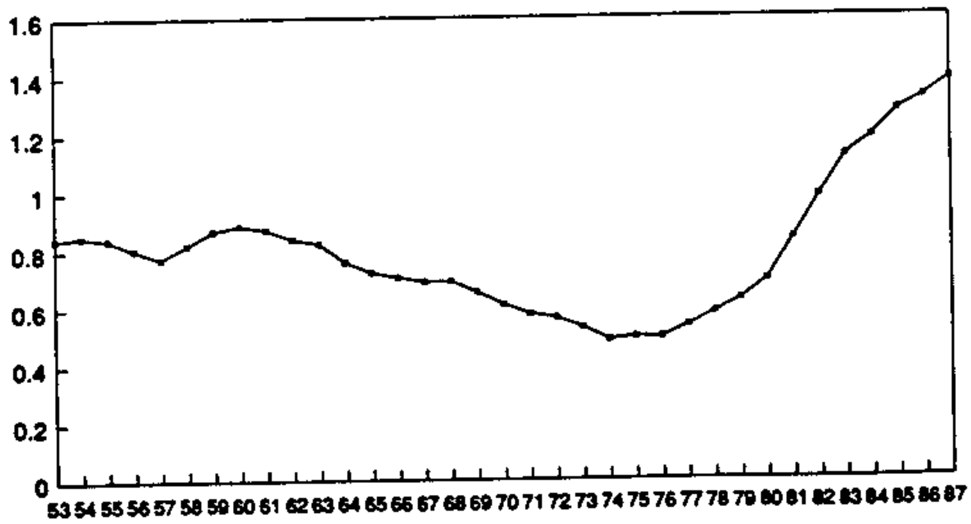


CHART 1 (cont'd)
Net public debt ratios

ITALY (1952-87)

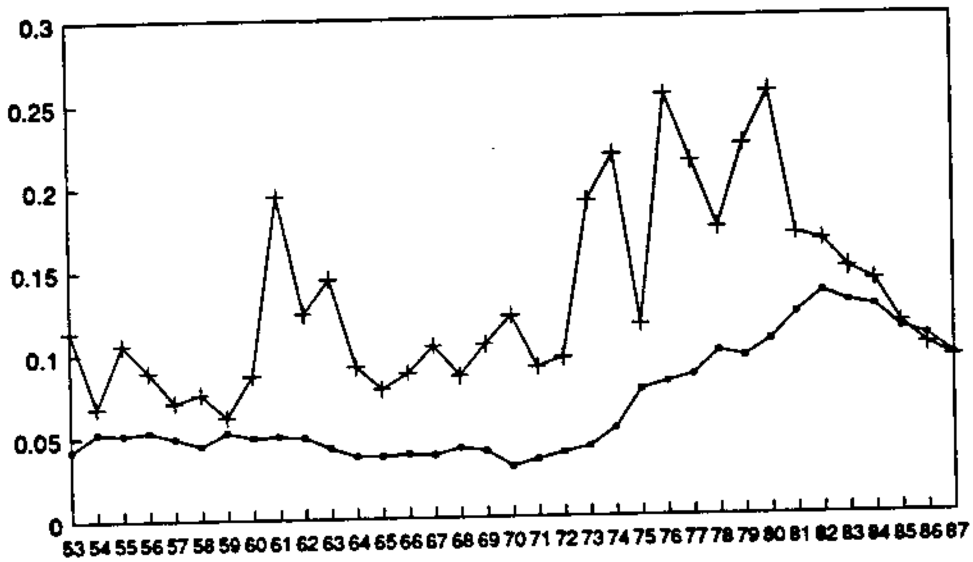


BELGIUM (1953-87)



—•— Net debt/Nat. income

CHART 1 (cont'd)
Debt service rate vs. income growth rate
ITALY (1953-87)



BELGIUM (1954-87)

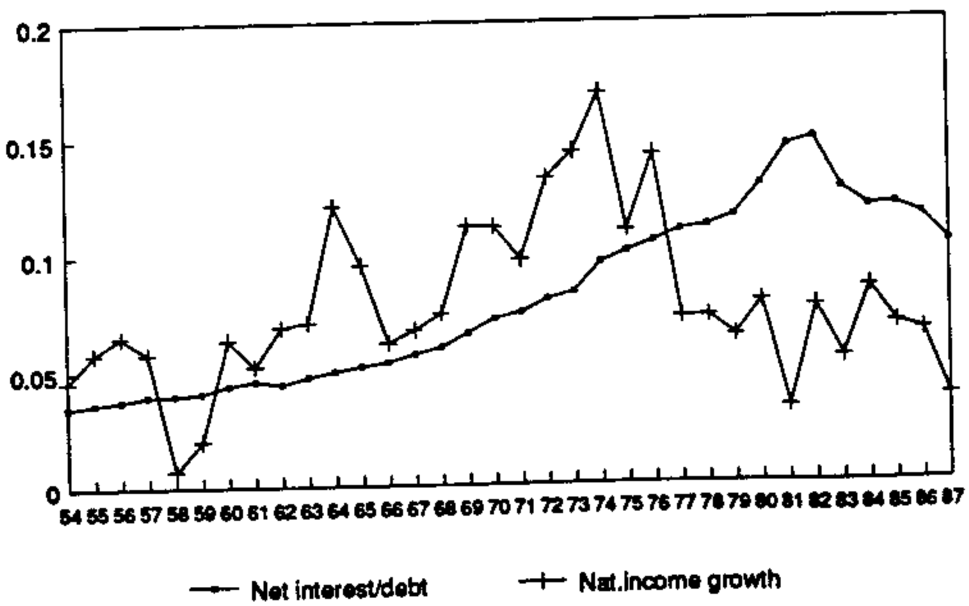
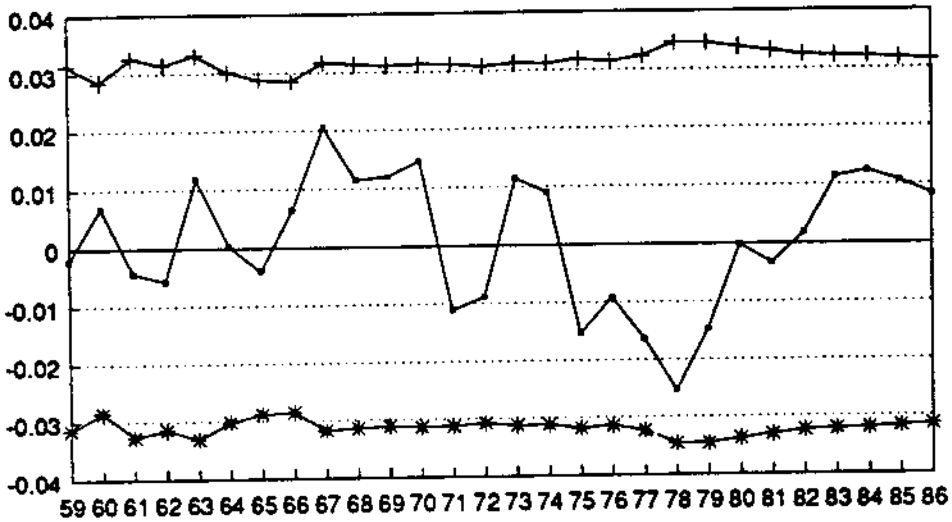


CHART 2

Cointegrating regression: ITALY

One-step residuals for Cn



One-step residuals for C

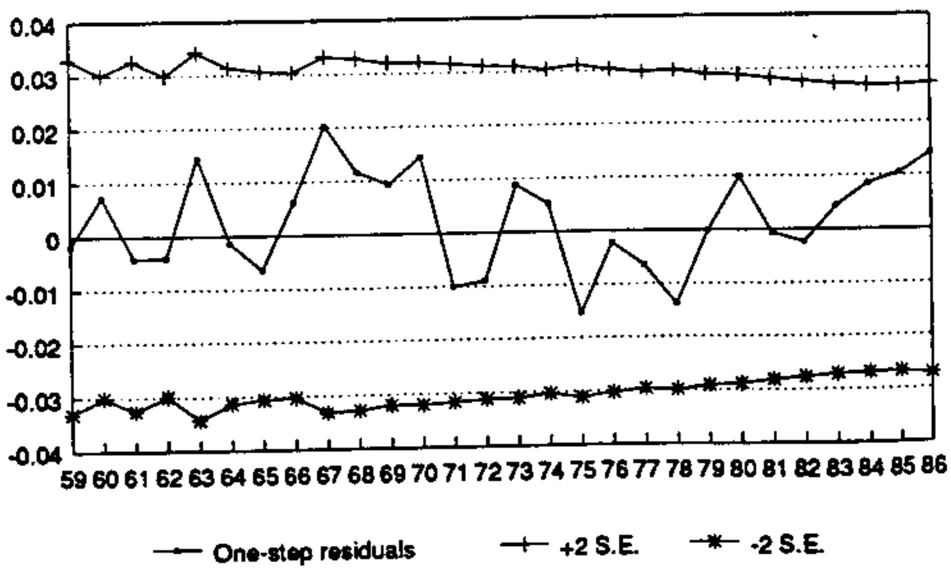
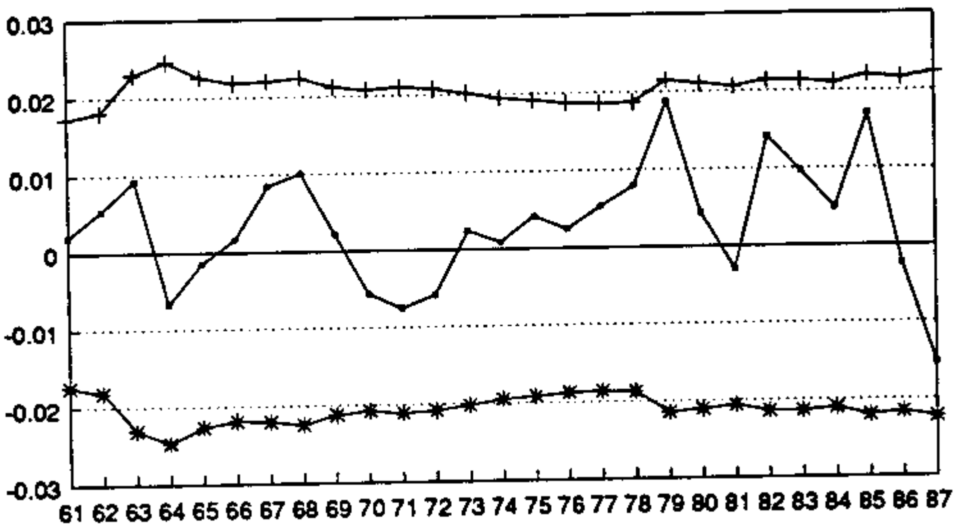


CHART 2 (cont'd)

BELGIUM

One-step residuals for Cn



One-step residuals for C

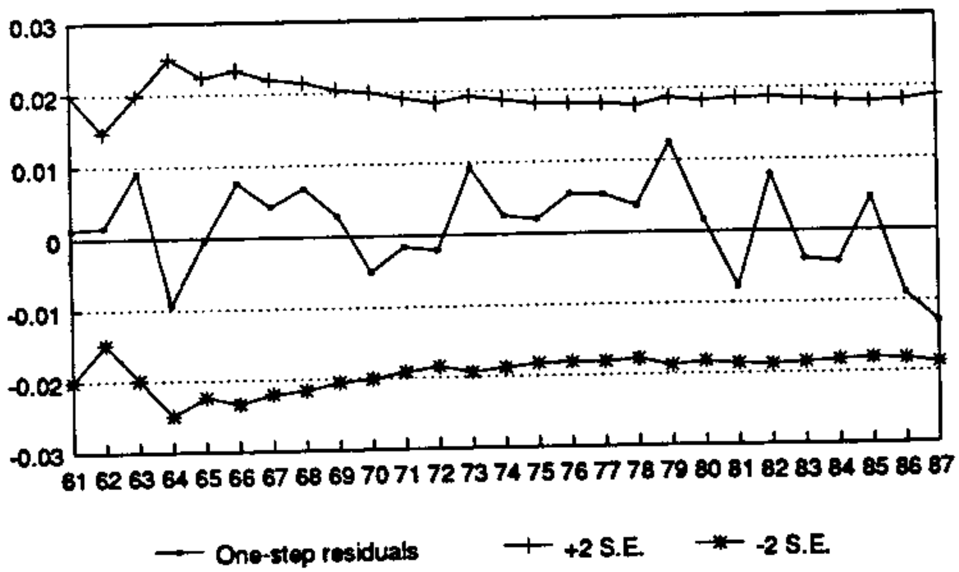
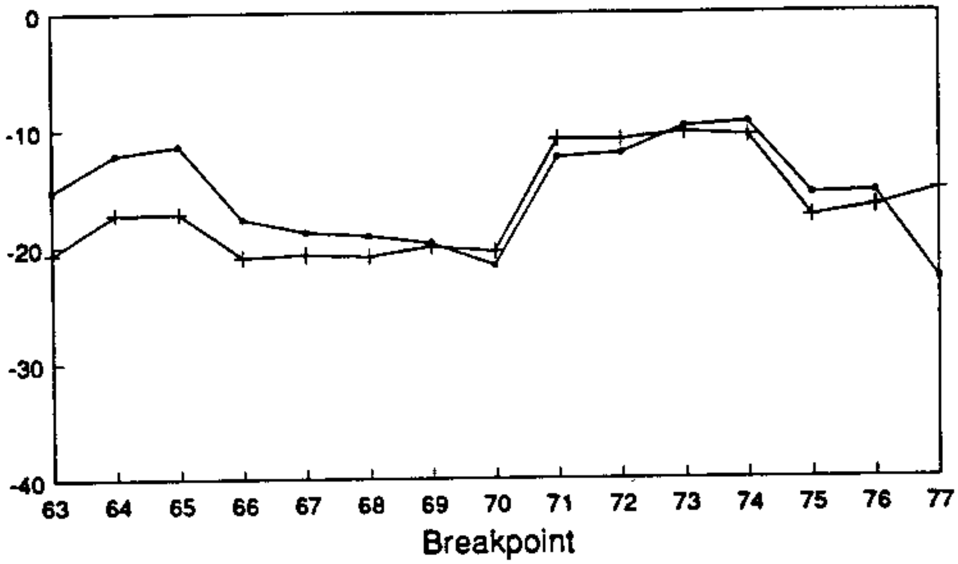


CHART 3
QUANDT'S RATIOS
Italy (1955-86)



Belgium (1957-87)

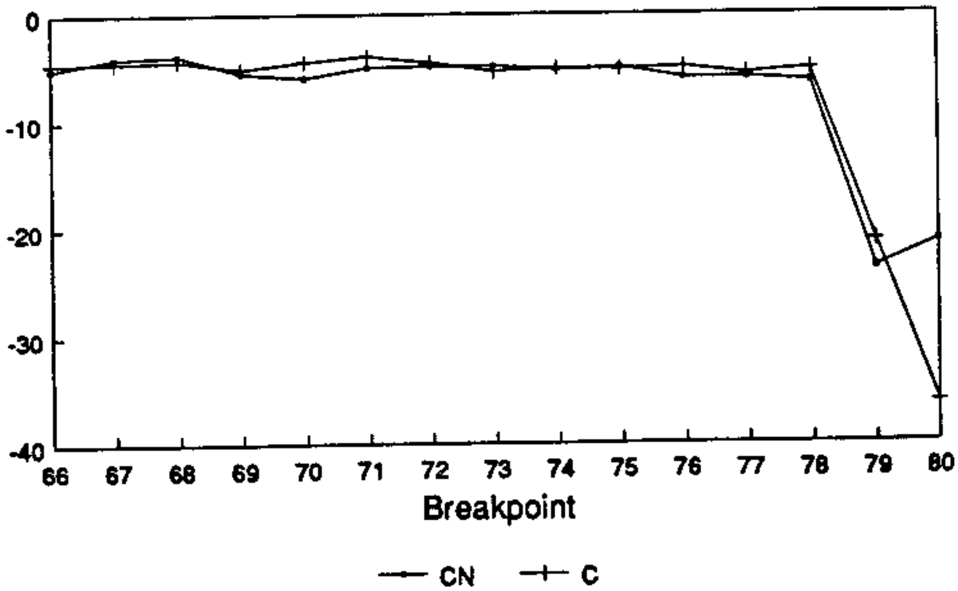
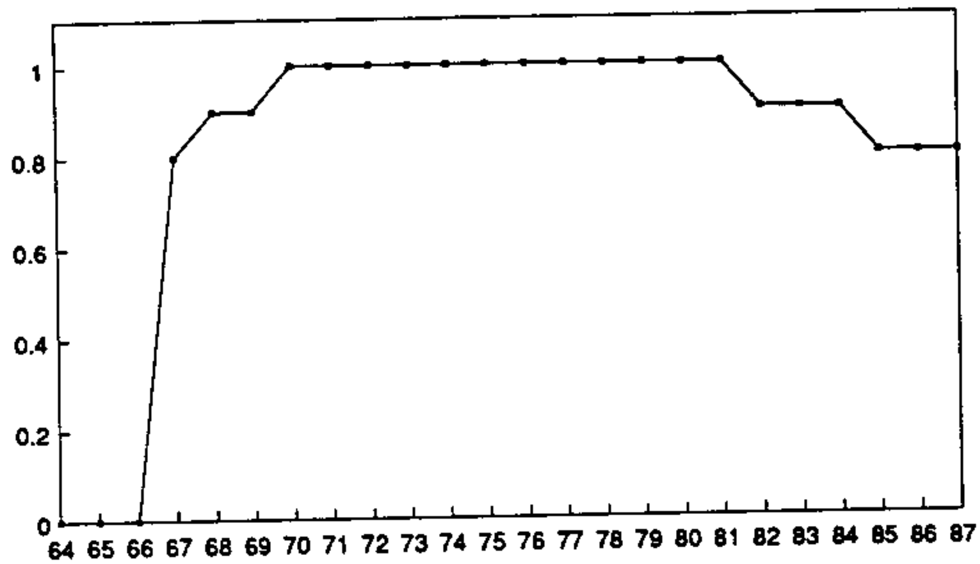


CHART 4
Recursive estimates of TD: ITALY

Cn



C

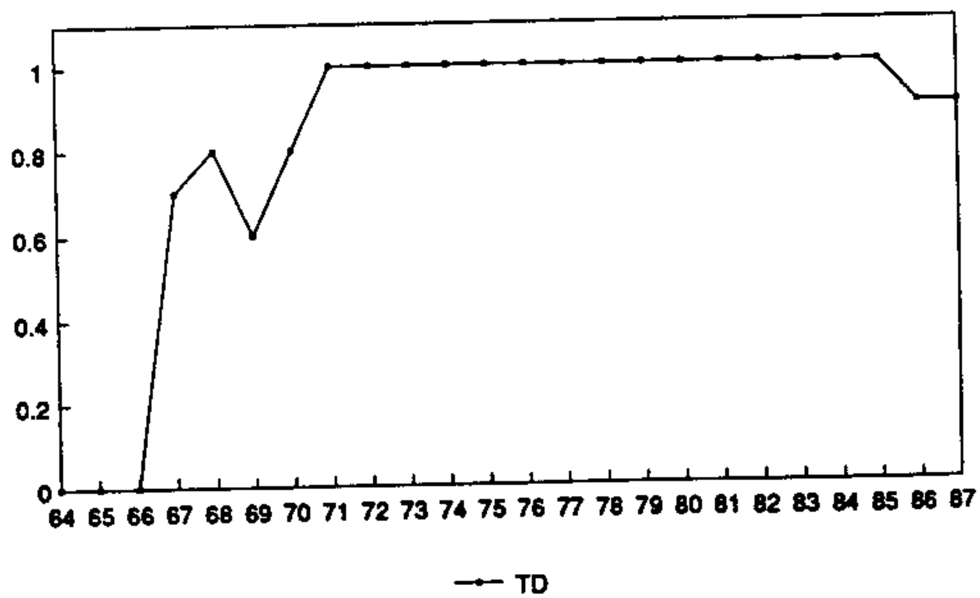
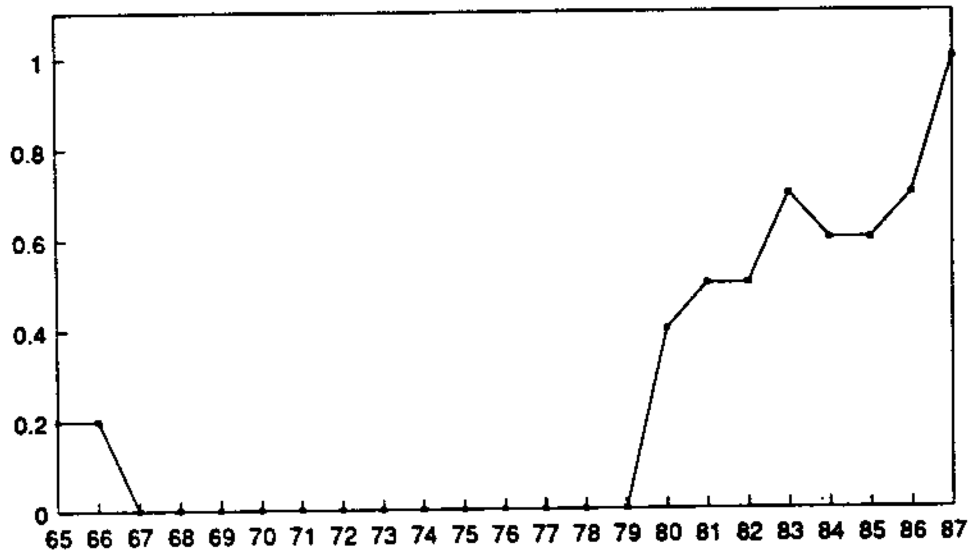


CHART 4 (cont'd)
Recursive estimates of TD: BELGIUM

Cn



C

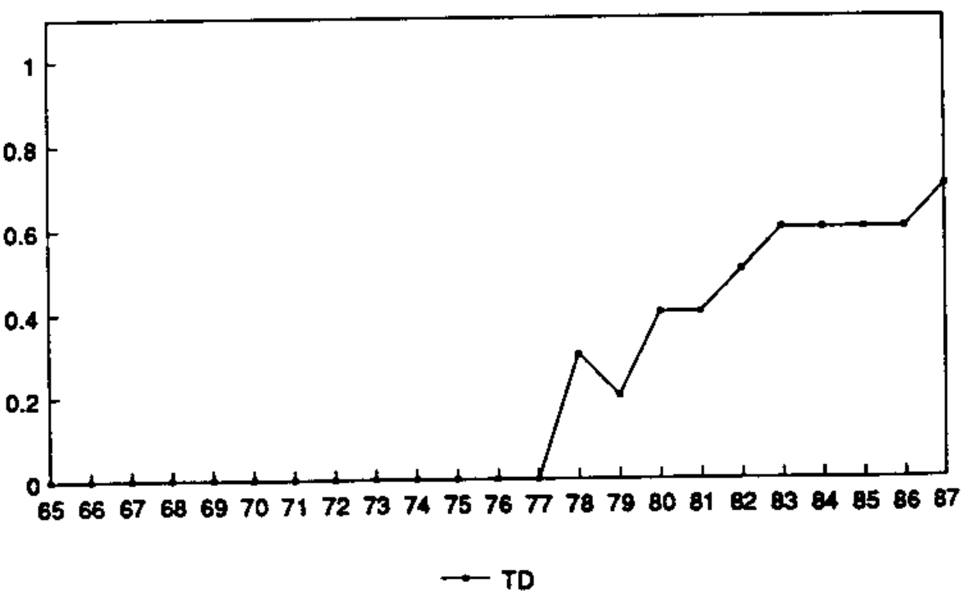
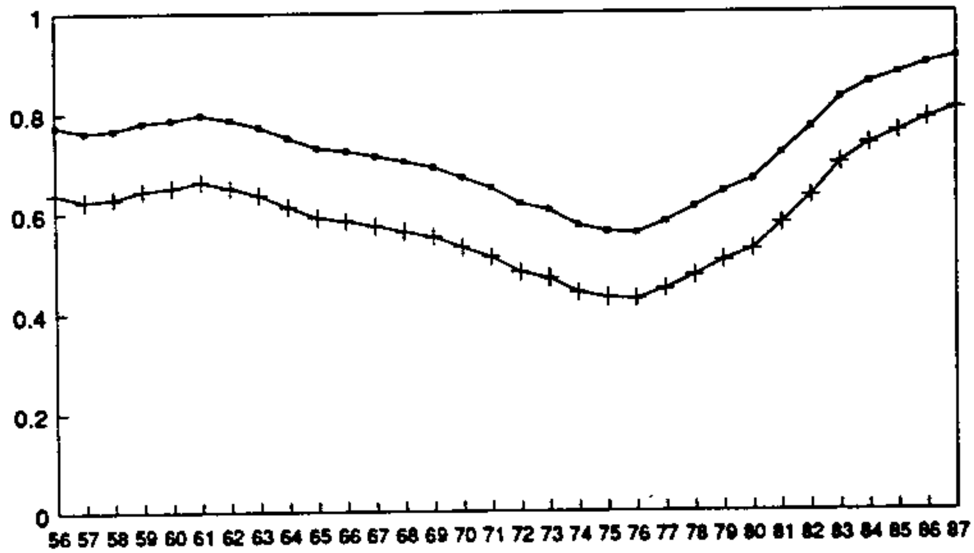
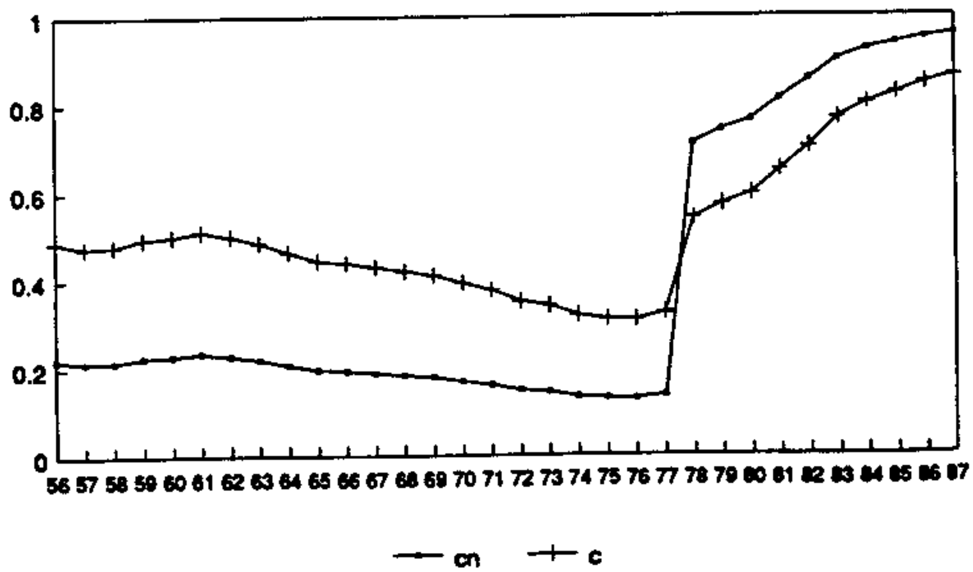


CHART 5
Variable TD model: BELGIUM (56-87)
td = 1-EXP[a*(b/y)]



Breakpoint: 1978
td = 1-EXP[a1*d1*(b/y)+a2*d2*(b/y)]



d1=dummy valued 1 in 57-78, 0 in 79-87
d2=dummy valued 0 in 57-78, 1 in 79-87