Faculty of Economics and Business Administration

2002-2003

FISCAL POLICY, GOVERNMENT DEBT AND PRIVATE CONSUMPTION.

Dissertation

Submitted at Ghent University,

to the Faculty of Economics and Business Administration,

in fulfilment of the requirements for the degree of Doctor in Economics

by

LORENZO POZZI

Thesis supervisor: Prof. Dr. F. Heylen

Fiscal Policy, Government Debt and Private Consumption.

by

Lorenzo Pozzi

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I would like to dedicate this dissertation to my father and my sister Lelia.

Lorenzo Pozzi,

7th of January, 2003.

Contents

Non-technical summary and conclusions.

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January 2003.

1. Orientation.

In the absence of a centralised European budget which redistributes income over member states if crises occur, countries joining the European Economic and Monetary Union (EMU) are de facto left with only national fiscal policy as an instrument to stabilise the economy. However, in the EMU, rules have been imposed on fiscal policy through the implementation of the Stability and Growth Pact. This pact has been established to keep countries from following unsustainable debt policies that could harm other member countries and / or put pressure on the European Central Bank. The literature offers no clear indication as to the effects that can be expected from respecting the rules imposed by the Stability and Growth Pact. Some authors suggest that the Pact limits the room for manoeuvre of fiscal policy makers, which undermines economic growth and intensifies recessions (see Eichengreen and Wyplosz 1998; Hughes Hallett and McAdam 1999). On the other hand, some argue that the Pact stimulates growth because it contributes to higher national savings (e.g. Loayza et al. 2000; Sarantis and Stewart 2001). Furthermore, it might make fiscal policy more effective because it induces a lower government debt (Nicoletti 1988; Sutherland 1997). It may also make fiscal consolidations, which remain an ongoing concern in many EMU countries, easier, for example by raising credibility.

In this context the following questions are of particular importance,

1. Fiscal policy being the only stabilisation instrument left in the hands of member states, it is very important that it is an effective instrument. Will it be ? And if it is (in)effective, why is that ? What is (are) the relevant channel(s) ? Throughout this dissertation, we define effectiveness as the extent to which changes in government expenditures and/or taxes affect output. As the title of this dissertation suggests, the effectiveness of fiscal policy is mainly measured by looking at its impact on private consumption.

2. Under what conditions will fiscal policy be effective ? Hemming et al. (2002) point to the potential role of a great number of elements: the exchange rate system, the openness of the economy, the business cycle situation, the monetary-fiscal policy mix, etc. Our emphasis is on the role of the government debt level in the effectiveness of fiscal policy. More precisely, we try to anwer the question whether a high government debt level increases or decreases the effectiveness of fiscal policy. Are high debt countries characterized by ineffective fiscal policy?

In section 2 we briefly discuss the main paradigms on fiscal policy and private consumption. In section 3 we discuss the specific research questions addressed in this dissertation and the general context of our reserach. In section 4 we discuss our main conclusions, their (scientific) relevance and their policy implications. In section 5 we present some ideas for future research.

2. Fiscal policy and private consumption: theoretical considerations.

In this section we present a brief overview of the main channels through which tax and government expenditure changes may affect the economy in Keynesian and neoclassical models. In Keynesian models (e.g. Fleming 1962, Mundell 1963) fiscal policy is typically effective. Both tax cuts and increases in government spending stimulate economic activity. The size of the output response is determined by a great number of elements : openness of the economy, exchange rate changes, reaction of monetary policy makers, interest sensitivity of investment and consumption, etc. By contrast, neoclassical models with optimising permanent income consumers imply the possibility of ineffective fiscal policy.

As to the effects of *tax changes*, an important starting point for our research is the so-called Ricardian Equivalence Theorem (Barro 1974). This theorem predicts that – for given government expenditures – tax changes will be neutral for the economy. Underlying this theorem are a number of strong assumptions like perfectly informed, rational and forward looking economic agents who discount lump-sum taxes, no precautionary savings, infinite horizons and perfect capital markets where no consumers are liquidity constrained (see Bernheim 1987, Seater 1993). If the theorem holds consumers 'pierce the government veil' and know that for instance a debt-financed tax cut must be repaid some time in the future. They save the amount of the tax decrease so that consumption remains unchanged. Dissaving by the government is fully compensated by the increased savings of the private sector. Barro (1978) acknowledges the importance of deviations from Ricardian Equivalence but notes that these deviations have indeterminate sign, leaving the Ricardian Equivalence hypothesis an appropriate benchmark to evaluate fiscal policy. If the underlying assumptions of the theorem are relaxed, tax changes will affect the economy. For example, if consumers with finite lives are assumed, tax reductions normally raise private consumption and economic activity through an increase in permanent income. Moreover, if taxes are distortionary, tax reductions will support the economy even more. Typically, however, the increase in private consumption and economic activity will be much smaller than in Keynesian models. The reason is that optimising neoclassical consumers smooth out any increase in their lifetime resources over all periods of their life.

As to *permanent* increases in *government spending* many neoclassical models predict that these generate a compensating one for one decrease in private consumption. Permanent increases in government spending need not be totally ineffective, however, when the induced reductions of private sector wealth stimulate labour supply for a given real wage. This is the case in so-called Real Business Cycle models (see e.g. Baxter and King 1993). Furthermore, changes in government spending may affect the marginal utility of private consumption and induce additional adjustments in the latter. This is typically the case in models that include the possibility of a so-called Edgeworth channel (Aschauer 1985, Karras 1994,...). If government spending and private consumption are Edgeworth substitutes (complements, independent) a cut in government spending will increase (decrease, not change) the marginal utility of private consumption and consumers will increase (decrease, not change) their expenditures.

The economy will also be affected if *spending cuts* are only *temporary*, since the impact of the spending cut on permanent income and thus on private consumption will be lower.

An interesting final element of neoclassical models based on the permanent income hypothesis is their stochastic implication. As shown by Hall (1978), permanent income consumers only react to 'news' about income, taxes and/or, if they (partially) incorporate the government budget constraint, to government expenditures[1].

What about the facts? Does Ricardian Equivalence hold, or is fiscal policy effective? Although existing empirical studies provide no clear-cut and unambiguous answer, our reading of the literature is that fiscal policy is indeed effective. A large majority of studies tends to reject strict Ricardian equivalence (see Loayza et al., 2000, Lopez et al. 2000, Sarantis and Stewart

[1] Note that Ricardian Equivalence is a logical extension of the permanent income hypothesis that includes the government sector.

2001, Doménech et al. 2000). As a very clear indication that fiscal policy matters, one could also refer to the fact that in most studies private consumption changes can be predicted on the basis of previously known information. In contrast to the stochastic implication of the permanent income hypothesis, private consumption changes are generally found to be 'excessively sensitive' to anticipated changes in (disposable) income and sometimes to anticipated changes in government expenditures. In the literature several explanations for this 'excess sensitivity' have been put forward. Typically these explanations come down to dropping one or more of the assumptions underlying extreme Ricardian and neoclassical views. 'Excess sensitivity' to income may be caused by liquidity constraints (Hayashi 1982, Flavin 1985, Campbell and Mankiw 1990), myopia (Flavin 1985), precaution (Carroll 1992, Barsky, et. al. 1986) or imperfect information (Pischke 1995, Demery and Duck 2000) and basically adds a Keynesian component to neoclassical models. 'Excess sensitivity' to government expenditures has generally been linked to the possibility of the Edgeworth channel that we have mentioned before. By allowing government spending to enter the utility function in a non-separable way, one can show that private consumption will not follow a random walk but will respond to anticipated changes in government expenditures.

As we have emphasised before, we pay special attention in this dissertation to the *influence of the level of government debt on fiscal policy effectiveness*. Interestingly, a number of existing studies on this issue suggest that high debt countries tend to have Ricardian characteristics (see Nicoletti 1988, 1992, Dalamagas 1993, 1994, Sutherland 1997, Slate et al. 1995,...). If the government debt level is high, the 'day of reckoning' may be imminent. Consumers may then be more aware of the future tax implications of the debt and may behave in a more Ricardian way.

3. Research questions and approach.

This dissertation contributes to the debate on neoclassical versus Keynesian theories, on the Ricardian Equivalence theorem and on the observation of 'excess sensitivity'. It contains four papers. In section 3.1. we discuss the specific research questions that we deal with in each paper. In section 3.2. we describe, evaluate and justify our theoretical and empirical approach. As will be seen, our approach is macroeconomic, with a focus on the analysis of private consumption within the Euler equation tradition.

3.1. Research questions.

In the orientation in section 1 we have raised two basic questions for our research. We address these questions in four papers. As far as the issue of the *effectiveness of fiscal policy* is concerned, all papers test for deviations from the neoclassical paradigm or from Ricardian Equivalence in one way or another. In paper 1, we estimate a consumption function for Belgium which nests two alternatives for Ricardian Equivalence that may cause fiscal policy to be effective. Consumers may be myopic with respect to (future) government activity and they may have a relatively large overall discount rate (reflecting a short horizon or a precautionary savings motive). Methodologically, in the estimation of the consumption function, we emphasize stationarity and small sample problems. In paper 2, we test for Belgium whether there is an effect of government expenditures on private consumption besides the standard negative effect that occurs when (Ricardian) consumers incorporate the intertemporal budget constraint of the government. More specifically, the question is raised whether private consumption and government consumption are (Edgeworth) substitutes, complements or independent. This issue is investigated in a methodological context, namely the possibility of obtaining unbiased estimates for the Edgeworth parameter and for the coefficient of relative risk aversion when estimating consumer Euler equations using a small sample. Given that, at the aggregate level, there is reason to be skeptical about an Edgeworth relationship, in paper 3 we provide an alternative explanation for the observed 'excess sensitivity' of private consumption to government expenditures. More precisely, as a theoretical innovation, we assume rational infinitely-lived forward-looking (Ricardian) consumers who incorporate the government budget constraint, but who are imperfectly informed about macroeconomic variables. Our consumers cannot distinguish idiosyncratic from aggregate components in their income and the taxes they pay. We test our model empirically for three high debt countries (Belgium, Italy, Greece). Paper 4 is also situated within the 'excess sensitivity' tradition (see e.g. Campbell and Mankiw 1990). In this paper the effectiveness of fiscal policy depends on the existence of (Keynesian) current income consumers versus (neoclassical) permanent income consumers. Current income consumers can explain why private consumption is excessively sensitive with respect to income. Effectiveness of fiscal policy results from the fact that consumption is then closely linked to disposable income and thus to taxes. Prominent explanations for the existence of current income consumers are liquidity constraints and myopia. We assess the relevance of these explanations for 19 OECD countries in the 1990s, using panel GMM estimation methods.

Regarding the second question raised in section 1, the *potential influence of the government debt level* on the effectiveness of fiscal policy is taken into account in papers 1 and 3 by estimating the derived consumption functions in high-debt countries (Belgium in papers 1 and 3, Italy and Greece in paper 3). In paper 4 we explicitly investigate whether the government debt ratio influences the effectiveness of fiscal policy by changing the fractions of current income versus permanent income consumers.

3.2 Basic approach.

As we have mentioned before, our focus in this dissertation is on private consumption. This choice is inspired first by the fact that private consumption expenditures constitute about 60% or more of GDP. As a consequence, they play a crucial role in any discussion on the effectiveness of fiscal policy. A second motivation is that testing for the effects of government deficits on other relevant variables like for instance the interest rate, the current account balance or the exchange rate would be more complicated because the expected effects may then depend on the assumptions concerning the openness and/or exchange rate system of the economy (see Seater 1993). In all four papers we follow the 'Euler equation' approach to private consumption. This means that consumption functions are estimated that are largely derived from first principles (and therefore include a limited set of regressors) and that test for specific departures from the permanent income hypothesis or Ricardian Equivalence. This approach differs from the more ad hoc estimation of reduced-form consumption or saving functions (containing a broad number of regressors) that allow a direct estimation of the effect of fiscal variables. The former approach has the advantage that it is based on utility maximization and rational expectations, whereas the latter approach has the advantage that it offers more flexibility. The ad hoc approach, however, more often suffers from misspecification and other methodological problems (see Seater 1993).

Some researchers may question our macroeconomic approach to analyse what are essentially microeconomic issues, i.e issues concerning individual consumer behaviour. Nevertheless, interest in consumption and saving is not confined to microeconomics but also concerns the behaviour of aggregate consumption. As we have mentioned, aggregate consumption is a major share of GDP. In order to understand the business cycle and how the business cycle can be influenced (e.g. for stabilisation purposes), it is necessary to understand the determinants of aggregate consumption. The study of the effects of fiscal policy studied in this dissertation must mainly be placed in this context.

Much of the recent macroeconomic literature on aggregate consumption works with 'representative agent' models, i.e aggregate consumption is treated as if it were generated by a single representative agent. While microeconomic studies have the advantage that they work with data that are appropriate from a theoretical perspective, the 'representative agent' model has some methodological advantages (see Deaton 1992, p.138-140): aggregation eliminates individual idiosyncracies (i.e consumer age, sex, family composition, race, education,...) and measurement errors so that estimation problems that result from them can largely be ignored. More importantly, given that there is a lack of household survey data that can be used to test the predictions of consumer theory, microeconomic studies may simply not be feasible. Moreover, conducting household surveys is not without problems. For instance, it may be difficult to construct a sample that is representative of the population. Moreover, although the questioning of individuals (see e.g. Allers et al. 1998) may be a more direct approach to investigate the behavioural assumptions underlying consumer theories than an approach based on aggregate data, it may not reveal much about true behaviour. As noted by Friedman (1953), people may behave in a certain way without being aware of the fact that they do. Deaton (1992) concludes that 'the econometric problems associated with testing the theory on microeconomic data are, if anything, more formidable than those associated with macroeconomic time series'.

Of course, the assumptions necessary to aggregate an individual's Euler equation into an aggregate Euler equation for a representative agent are, to say the least, very demanding. Besides the usual restrictions that have to be placed on the individual preferences (i.e the

Gorman form), there are additional issues that are specific to intertemporal choice under uncertainty. As shown by Grossman and Shiller (1982), for the Euler equations to aggregate perfectly it is also necessary that (a) people live forever (or stated differently, are intergenerationally linked through an altruistically motivated operative bequest motive) and that (b) people all know and use the same aggregate information. Under these unrealistic conditions, aggregation is straightforward and thus irrelevant [2].

In this dissertation the strict representative agent model is used only in paper 2. The focus in this paper lies on a methodological issue, namely the finding of poor estimates for the coefficient of relative risk aversion because of the use of a small sample. As acknowledged in the paper, this problem may also be due to the underlying model or because of aggregation issues. In paper 1, following Blanchard (1985), a model for aggregate consumption is considered where the implicit assumption is that individual consumers are different from each other because they are born in a different period. In paper 3 it is assumed that consumers are different because they have different information sets. In paper 4 we start from a typical empirical specification which treats aggregate consumption as the result of the consumption of two different consumer groups: permanent income consumers and current income consumers. In that sense the papers in this dissertation are based on less restrictive assumptions than those on which 'representative agent' models are typically built and do, as a result, take into account some of the empirical deviations from the permanent income hypothesis encountered in the literature (e.g. excess sensitivity, excess smoothness, consumers with a discount rate that is higher than the real interest rate,...). Methodologically, they maintain the methodological advantages associated with the 'representative agent' aggregate consumption approach which we mentioned above.

[2] Note that if aggregation is deemed unimportant, it is not possible to view the aggregation itself as a potential explanation for the failure of the strict permanent income hypothesis on aggregate data, i.e as a source of the empirical finding of 'excess sensitivity' of private consumption to income or as a source of other empirical deviations from the strict permanent income hypothesis.

4. Conclusions, relevance and (policy) implications.

In this section we discuss the main conclusions of each paper and their implications for (the effectiveness of) fiscal policy. We also emphasize the relevance of each contribution. We then formulate a general conclusion and we discuss the policy implications of this conclusion for EMU countries.

Our main conclusion in paper 1 is that we cannot reject for a high-debt country like Belgium that fiscal policy actions through taxes are effective. The main reason is that consumers appear to have a relatively high overall discount rate. This may be because consumers have a planning horizon that is shorter than the government's horizon or because they have a precautionary savings motive. However, since the results also suggest that consumers strongly take into account future government activity, the effectiveness of tax changes may be lower than if consumers were more myopic with respect to future government activity. This result is in line with previous findings in the literature which suggest that the Ricardian Equivalence hypothesis may hold only partially (see section 2). The main contribution of the paper to the literature lies in the use of a model in which two alternatives to Ricardian Equivalence are incorporated, instead of only one.

Neither in paper 1 nor in paper 2 do we find evidence that an Edgeworth channel for fiscal policy is present in Belgium. A methodological implication of paper 2 is that the common practice of taking a poor estimate for the coefficient of relative risk aversion for granted (in estimated consumer Euler equations) and nevertheless draw conclusions from the other estimated parameters in the model (see e.g. Holman 1998) should be avoided.

In paper 3 we show that the empirical observation that private consumption is excessively sensitive to government expenditures may be the result of the presence of imperfectly informed consumers who incorporate the intertemporal budget constraint of the government. We find that this model with imperfectly informed Ricardian consumers seems to describe the data rather well for high-debt countries like Italy and Greece during periods of high debt accumulation. The model does not work well for Belgium however. The implication is that, at least for Italy and Greece, tax changes may be relatively ineffective and anticipated and unanticipated government expenditure increases seem to have a negative effect on aggregate consumption. As in paper 1 we thus find Ricardian characteristics in high-debt countries, though strict Ricardian Equivalence is again rejected. The main contribution of the paper is that it provides an alternative to the standard interpretation that the excess sensitivity of private consumption to government expenditures is caused by an Edgeworth channel. The latter type of relationship may be questioned at the aggregate level.

Our results in paper 4 suggest that liquidity constraints are endogenous to government debt (after controlling for the effect of financial liberalization). This has important implications for the effectiveness of fiscal policy. On the one hand, fiscal policy is more effective (Keynesian) if the number of current-income liquidity constrained consumers is higher. A debt-financed tax cut, for instance, will then increase consumption because liquidity constrained consumers approach their desired optimal consumption path. On the other hand, as shown theoretically by Yotzusuka (1987), if a higher government debt reinforces liquidity constraints, a debt-financed tax cut is not only good news for liquidity constrained consumers (who can spend more). It is also bad news for initially unconstrained consumers who cannot borrow anymore (or can only borrow less) because banks tighten credit conditions. These consumers may then have to postpone, say, the purchase of a home or a durable, which undermines the effectiveness of fiscal policy. The net effect of fiscal policy on private consumption is unclear. Thus strict

Ricardian Equivalence is again rejected, now because liquidity constraints seem to be relevant to explain aggregate consumption. The endogeneity of liquidity constraints to government debt may (partly) re-establish a Ricardian result however through the forward-looking behaviour of the banking sector, not of the consumers. The findings in paper 4 also reveal a weakness in recent studies investigating the effects of fiscal policy on private consumption and savings (e.g. Evans and Karras 1998, Perotti 1999, Lopez et al. 2000). All these studies assume the existence of two or three groups of consumers, the fractions of which are considered constant (e.g. permanent income consumers, liquidity constrained consumers, etc.). Our results challenge this assumption. The stance of fiscal policy itself and the government debt ratio may change these fractions.

Let us then return to the two basic questions that we raised in section 1 of this chapter. As a general conclusion with respect to the first question we can state that we do not doubt the effectiveness of fiscal policy. Relevant channels of influence that show up from our research are finite horizons, imperfect information on aggregate variables, the existence of liquidity constraints and precautionary savings. The Edgeworth channel seems to be irrelevant.

As to question 2, we do find some Ricardian characteristics in the high debt countries considered in papers 1 and 3 and in our panel of countries studied in paper 4. There seems to be some evidence (in accordance with some of the literature) that the effectiveness of fiscal policy is negatively related to government debt level [3].

[3] We note here that, in papers 1 and 3, we test our model using data for high-debt countries (Belgium in papers 1 and 3, Greece and Italy in paper 3) because some of the literature suggests that Ricardian Equivalence is more likely to hold in these countries (see Nicoletti 1988, Dalamagas 1993, 1994). We tend to conclude that the high-debt economies considered have Ricardian characteristics. This conclusion is weakened because it does not hold for Belgium in paper 3. Moreover, this conclusion does not exclude the possibility that low debt countries also have Ricardian characteristics. Given the literature it seems more logical however that the benefits of piercing the government veil are larger for consumers in countries with a problematic government debt. The estimation for the US (a country with a normal debt situation) of a model similar to the one considered in paper 1 does not reveal any Ricardian characteristics of US consumers (see Graham and Himarios 1991).

Given the current worldwide economic downturn governments in some EMU countries could be tempted to abandon the requirements of the Stability Pact so as to increase the possibilities of using the fiscal policy instrument to fight the economic crisis. Given the controversy in the literature on the effects of the Stability Pact (see section 1), it is difficult to judge whether this would be a sensible decision. Given some of the conclusions of this dissertation however, we feel that governments should be cautious to abandon fiscal rules. Even if our results do not suggest that fiscal policy is ineffective, in the long run the accumulation of debt that results from it may change the reaction of economic agents to fiscal policy and diminish its effectiveness. More specifically, some of our results seem to imply that economic agents in high-debt countries and financial institutions tend to behave in a Ricardian way, i.e they seem to have some 'awareness' that governments that stimulate the economy by raising debt will have to increase taxes some time in the future to repay this accumulated debt. Therefore, in response to a tax decrease consumption may not increase as much as it would under low-debt circumstances.

As noted in a recent literature review on the effectiveness of fiscal policy by Hemming et al. (2000), the theoretical literature suggests that fiscal multipliers are, ceteris paribus, likely to be 1) smaller when households are more Ricardian and 2) higher if the government debt is low. All these results point towards high government debt levels decreasing fiscal multipliers. High government debts may make it harder for governments to fight recessions. Given that, as noted by Perotti (2002) and Hemming et al. (2002), empirical estimates of fiscal multipliers in OECD countries are overwhelmingly small already, high government debts might just make the multipliers even smaller. In this context, it is reassuring that the European Commission has stated recently that, apart from its focus on deficits, it will devote more attention to the (reduction of) countries' debt levels. Low-debt countries without a deficit will receive more flexibility in using the fiscal policy instrument. In our opinion, another necessary measure would be to make fines for offenders more enforceable. A current problem, given that recessions tend to be correlated across countries, is that it may be hard to find a majority of countries willing to punish an offending country.

5. *Directions for future research.*

In this section we describe a number of ideas as to how the research conducted in this dissertation in particular and in the literature on private consumption, fiscal policy and government debt in general could be extended.

In paper 1 of this dissertation we consider the issue of tax discounting, which depends on the degree of consumer myopia with respect to government activity. The latter parameter is assumed to be time-invariant. In line with the time-varying tax discounting set-up by Nicoletti (1992), it might be interesting to assume that this parameter is time-varying and dependent on the government debt level. Alternatively, it would be interesting to estimate this parameter for different countries and link this to the government debt level in these countries. Panel data techniques (see Lopez et al. 2000, see paper 4) could then be especially useful to alleviate some of the mentioned estimation problems by increasing sample sizes.

The testing of fiscal policy effects on private consumption in general and Ricardian Equivalence in particular could make use of richer empirical models than the one considered in paper 1. It might be interesting to consider models with different consumer types (see Lopez et al. 2000) : (in)finitely lived consumers, liquidity constrained consumers (see paper 4 in this dissertation), etc.

The type of myopia considered in paper 1 takes the form of consumers not taking into account current and future government activity. It seems unlikely however that consumers will completely ignore taxes since they affect disposable income and consumers receive their income net of taxes. Considering alternative, more general, forms of myopia (e.g. consumers may be myopic in the sense of expecting constant future tax rates) may also be an interesting extension of chapter 1.

In paper 2 we find that, when estimating consumer Euler equations using a small sample, the estimate for the coefficient of relative risk aversion can be seriously biased. It would be interesting to investigate how the biases are affected and how fast they disappear when more data are used (e.g quarterly data). Given the lack of quarterly data for Belgium, the analysis could be conducted for other countries. Another approach might be to generate data out of a completely specified and calibrated economic model.

The imperfect information model considered in paper 3 is derived under the certainty equivalence assumption. It might be interesting to extend it to allow for utility of the constant relative risk aversion type instead of quadratic utility. Given the analytical difficulties of this approach, an alternative might be to allow the discount rate of consumers to differ from that of the government (as in paper 1 of this dissertation). This can be reconciled with precaution and also with finite horizons. Different discount rates for consumers and the government imply that consumers do not completely incorporate the intertemporal budget constraint of the government. Thus there is the possibility of excess sensitivity of private consumption not only to pre-tax income and government expenditures but also to taxes.

As a further extension of paper 3, it might be interesting to empirically determine good approximations for the processes for income and government expenditures so that a more direct test of the theory on the data is possible.

In paper 4 we briefly address the link between the government debt, uncertainty and precaution and private consumption. High government debts may increase the uncertainty about future consumption. This will result in a steeper consumption path. The effect of government debt on excess sensitivity in paper 4 is only interpreted in terms of the liquidity constraints channel. Excess sensitivity of consumption to current income may also be affected by the size of the government debt because of the uncertainty about future income or consumption that it creates. First, as noted by Carroll (1992), if current income is a good predictor of the future variability of consumption, excess sensitivity may arise. More interestingly, as noted by Barsky et al. (1986), excess sensitivity of consumption to current income may also arise if taxes are positively related to income and uncertainty about future income is substantial. In this framework, a tax cut leads to increased consumption because it provides certain wealth while the associated future tax increase is contingent upon future (uncertain) income. Barsky et al. further show that the higher the risk involved in future income, the stronger the propensity to consume out of a current tax reduction. In this framework also, a higher government debt may lead to more excess sensitivity.

To give a correct interpretation to the positive statistical relationship between government debt and the excess sensitivity of private consumption to current income, it may be necessary to discern the precaution channel from the liquidity constraints channel.

We also mention directions for future research that do not stem directly from the theoretical or empirical issues discussed in this thesis. These directions are part of a literature that has been largely disregarded in this thesis but nevertheless constitutes an important part of the literature on fiscal policy, government debt and private consumption. First, there is a stream of literature that emphasizes that the effects of fiscal policy may be completely reversed depending on the government debt level (e.g. see Sutherland 1997, Perotti 1999). For instance, tax and/or government expenditure changes may be Keynesian at low debt levels and non-Keynesian (but still effective) at high debt levels. There is room for a wide range of possible empirical contributions to this literature (e.g. panel data studies). Second, estimating empirical consumption or savings functions (see e.g. Sarantis and Stewart 2001) and relating the coefficients of these functions (e.g. the coefficient on the government deficit) to the government debt may be a possible strategy for future research. To the extent that the variables in these consumption or savings functions are expressed in levels, a major advantage of this type of studies is that it is possible to take into account potential long-run relationships between variables (consumption-income,...) through (panel) cointegration techniques. Related to this is the issue of testing the long-run *assumptions* of fiscal policy models (e.g intertemporal budget constraints of consumers and government, see Becker 1997). We also mention the problem of incorporating long-run assumptions of fiscal policy models into Euler equation studies. In the literature some studies use error correction terms as instruments directly in the estimation procedure of the Euler equation without testing their validity. Other studies simply neglect the long run. It would be interesting to investigate in how far these different approaches affect the conclusions reached.

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Tax discounting in a high debt economy.

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Abstract.

This paper estimates a consumption function for Belgium which allows for government debt discounting and for the overall discounting of the future (reflecting the consumers' planning horizon or precautionary savings). It also allows for substitutability or complementarity effects from government expenditures. We use bootstrapped distributions for inference since the instrumental variables estimators used may have non-standard distributions. This procedure also helps to tackle potential endogeneity and sample size problems. Results suggest that consumers do take into account (future) government activity. Ricardian Equivalence is rejected however, since we cannot reject a relatively short planning horizon or a precautionary savings motive for the consumers.

JEL classifications : E21, E62.

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1. Introduction.

To assess the effectiveness of fiscal policy the debate on the Ricardian Equivalence proposition is of particular relevance. The Ricardian Equivalence theorem (Barro 1974) basically states that rational consumers 'pierce the government veil'. They understand that the government is bound to satisfy an intertemporal budget constraint. This implies that if the government, for instance, lowers taxes today it will have to raise them tomorrow, given a stream of government expenditures. Under a set of assumptions (see Bernheim 1987 and Seater 1993), among which are infinite planning horizons, no liquidity constraints and no precautionary savings motives for consumers, rational agents do not alter their consumption when confronted with a tax decrease. Instead, they save more because they expect future tax increases. Private savings thus rise and make up for the rise in the government deficit so that the total savings in the economy do not change. Interest rates and aggregate demand are unaffected.

In this paper we consider a model with two alternatives to Ricardian Equivalence. First, consumers may have a finite horizon (see Blanchard 1985 and Evans 1988) and may thus be myopic with respect to all future variables. In the model this is reflected by a consumer's discount rate that exceeds the real interest rate. Alternatively, the difference between the consumer's discount rate and the real interest rate can be interpreted as a risk premium stemming from a precautionary savings motive (Hayashi 1982, Muellbauer and Lattimore 1995). Second, consumers may be myopic solely with respect to government activity (see Graham and Himarios 1991). Both alternatives imply that the incorporation of the intertemporal budget constraint of the government and thus the incorporation of the future tax implications of current debt will be incomplete. Only if consumers take into account government activity and have infinite horizons (or do not have a precautionary savings motive), government bonds are not perceived as wealth and Ricardian Equivalence holds. Contrary to existing research, we thus simultaneously consider two alternatives in a consistent way when determining the degree of tax discounting.

Deviations from the Ricardian benchmark may be linked to specific characteristics of an economy. In this paper we are particularly interested in whether the debt situation in a country affects the degree of tax discounting. The literature suggests that consumers in high debt countries are more aware of the future tax implications of current debt. Nicoletti (1988) finds evidence of tax discounting for high-debt countries like Belgium and Italy but not for the UK and Germany. Strong evidence in favour of this hypothesis is also given by Nicoletti (1992) for Belgium, where tax discounting seems to be time-varying and increasing with the debt ratio.
We test our model for Belgium where the government debt has exceeded 100% of GDP in most of the 1980s and 1990s making it a country for which, a priori, we would expect to find Ricardian characteristics.

Anticipating the results, our estimates for the parameter that captures the possible incorporation of government activity seem to imply that Belgian consumers are quite aware of (future) government activity. We suggest that this may be related to the high government debt level in Belgium. Complete Ricardian Equivalence is rejected, however, since the estimates of the overall discount rate of consumers are such that we cannot dismiss that the consumers' horizon is shorter than that of the government or, alternatively, that consumers have a precautionary savings motive.

An additional feature of our model is that it allows to test whether private consumption and government expenditures are Edgeworth substitutes, complements or independent (see Aschauer 1985, Ni 1995). This captures an additional potential effect of fiscal policy on private consumption. We find no evidence to support or reject this possibility however.

From the above discussion it should be clear that different empirical results on Ricardian Equivalence across countries can probably to a large extent be attributed to differences in the economic conditions (e.g the debt level) of these countries. Note however that empirical work on tax discounting also finds contradictory results within countries. For the US, for instance, the tax discounting hypothesis has been rejected by among others Feldstein (1982), Nicoletti (1988) and Graham and Himarios (1991, 1996). It has not been rejected by among others Kormendi (1983), Aschauer (1985) and Evans (1988).

Seater (1993) notes that underlying the disparity in empirical results may be methodological problems like simultaneity and non-stationarity. In this paper we especially focus on the issue of stationarity. In the literature, consumption functions similar to the one we derive are estimated without taking into account that the estimates may have a non-standard asymptotic distribution. The reason is that non-stationary variables enter the equation in levels. It makes inference problematic. We conduct a (block) bootstrap procedure to try to avoid this pitfall. Inference based on small sample distributions also makes sense given that our estimations must be conducted with a small number of observations.

The paper is organised as follows. In section 2 we discuss the model leading to the consumption function to be estimated. In section 3 we tackle data and methodological problems. In section 4 we present our empirical results and some robustness tests. Section 5 concludes.

2. The Model.

In this section we present a model in which consumers can be myopic with respect to all future variables or, alternatively, can have a precautionary savings motive. Consumers may also be myopic solely with respect to future government activity. Ricardian Equivalence shows up as a special case. The model leads to a consumption function which is then estimated in the next section.

We start from the standard discrete time representation of Blanchard's (1985) aggregate consumption function (see Evans 1988; Graham and Himarios 1996; Haug 1996), written in terms of 'effective' rather than private consumption,

$$(1)\quad c_t^* = \beta\left[\sum_{j=0}^{\infty}(1+r+\rho)^{-j}E_t(y_{t+j}-t_{t+j}+\lambda g_{t+j})+w_t+b_t\right]+v_t$$

where c_t^* is total or 'effective' consumption, y_t is pre-tax labour income, t_t are net taxes, g_t are government expenditures, w_t is financial wealth (excluding government bonds) measured at the beginning of period t and b_t are one-period government bonds measured at the beginning of period t. All variables are in real per capita terms. E_t is the expectations operator conditional on information available at time t. The term between square brackets in (1) is permanent income or wealth and the parameter β is the marginal propensity to consume out of permanent income. The term v_t reflects transitory consumption or a preference shock [1].

The consumers' discount rate is given by r+ρ, where ρ is a mark-up over the real interest rate r. The latter is assumed to be time-invariant. Blanchard (1985) justifies that the discount rate differs from the real interest rate by the finite horizon assumption. The higher ρ, the shorter is the horizon of the consumers, so that it will take only a few periods before future variables are irrelevant for today's consumption. When ρ=0, the consumers' planning horizon is infinite as is the government's. An alternative interpretation for ρ>0 is that consumers have a precautionary savings motive when they face uncertainty. To calculate human wealth they discount expected uncertain future income and taxes at a rate higher than the interest rate so that ρ can be thought of as a risk premium (see Hayashi 1982; Muellbauer and Lattimore 1995).

The parameter λ is related to 'effective' consumption, which is defined as a linear combination of private consumption and government spending,

(2) $c_t^* = c_t + \lambda g_t$

where c_t is private consumption. If private consumption and government expenditures are Edgeworth substitutes (complements, independent) we have that $\lambda > 0$ ($\lambda<0$, $\lambda=0$) (see Feldstein 1982, Kormendi 1983, Aschauer 1985, Karras 1994)[2]. The term $\lambda\sum_{j=0}^{\infty} (1+r+\rho)^{-j} E_t g_{t+j}$ in (1)

[1] The microeconomic foundations for this type of consumption function can be found in Blanchard (1985). Evans (1988) employs a discrete time, stochastic representation of Blanchard's model.

[2] Take the total differential of (2) and set equal to zero. This leads to $dc_t / dg_t = -\lambda$.

reflects the potential wealth effects related to whether private consumption and government expenditures are substitutes or complements (see Graham and Himarios 1991, Karras 1994). There will be a positive wealth effect if $\lambda>0$, for instance, because government expenditures diminish the need for certain private expenditures and thus leave the consumers with more resources. If $\lambda=0$ this implies that private consumption decisions are not affected by changes in expected government expenditures other than the ones coinciding with changes in expected taxes.

Introducing a parameter of myopia θ in (1) to capture the possibility that consumers are myopic with respect to government activity, but not with respect to their own labour income, we obtain,

$$(3)\quad c_t^* = \beta\left[\sum_{j=0}^{\infty}(1+r+\rho)^{-j}E_t(y_{t+j}-\theta(t_{t+j}-\lambda g_{t+j}))+w_t+b_t\right]+v_t$$

with $0 \leq \theta \leq 1$ (see Graham and Himarios 1991). The parameter θ reflects the extent to which consumers perceive future government expenditures and net taxes as affecting their wealth. If θ = 1 consumers take into account (future) government activity and equation (3) reduces to equation (1). The degree of discounting of government activity then only depends on ρ, the overall weight the consumers gives to the future. If $\theta = 0$ consumers are completely myopic with respect to government activity, whatever the value of ρ (and thus whatever the length of their planning horizon or the strength of their precautionary savings motive) [3].

[3] Graham and Himarios (1991) justify the type of myopia captured by θ by referring to Kormendi (1983) who notes that the standard permanent income equation, as represented by (1), implies quite an asymmetry with respect to consumer behaviour regarding fiscal policy. Consumers are assumed to be perfectly forward looking with respect to future income and taxes. Government debt b_t is nevertheless included as a part of private wealth, the implicit assumption being that consumers are too myopic to take into account the future tax implications of current debt.

Equation (3) has features that may cause some concern. First, note that $\theta=0$ is a limit case. Finding complete myopia may be rather unlikely because consumers will, at least to some extent, be aware that taxes affect disposable income. Note also that other forms of myopia may be possible which are not captured by the current analysis[4]. Second, if $\lambda\neq0$ this may have implications for the interpretation of θ. Certain forms of irrationality may appear if $\lambda\neq0$. One could imagine cases where even $\theta=1$ corresponds to a form of myopia[5]. As an extreme example, suppose $\theta = 1$ and $\lambda = 1$. In that case, if the budget is always balanced, the consumption choice defined by (3) disregards taxation. Both concerns may not be that problematic, however. Anticipating our results, we cannot reject that $\theta = 1$ *and* $\lambda = 0$. In other words, we cannot reject one aspect of rational and forward-looking behaviour (consumers take into account expected taxes). Neither can we reject that, conditional on expected taxes, private consumption is independent from expected government expenditures.

The government budget constraint in period t is given by,

$$(4) \quad (g_t - t_t + b_t)(1+r) = b_{t+1}$$

We use (4) into (3) to obtain the optimal consumption rule (see Appendix A1),

$$(5) \quad c_t^* = \beta\left[\sum_{j=0}^{\infty}(1+r+\rho)^{-j}E_t\left(y_{t+j} - \theta(1-\lambda)g_{t+j} + \theta\rho(1+r+\rho)^{-1}(1+r)^{-1}b_{t+1+j}\right) + w_t + (1-\theta)b_t\right] + v_t$$

[4] E.g. consumers may be myopic in the sense that they assume that future tax rates are unchanged.

Let us first consider the infinite horizon / no precaution case, i.e. $\rho = 0$. If, in addition, $\theta = 1$ consumers have the same infinite horizon as the government and completely internalise the government budget constraint. There is a full incorporation of future taxes and Ricardian Equivalence holds. In that case, as can be seen from (5), private consumption depends only on w_t and government bonds b_t are not considered as a part of wealth. If, on the other hand, $\theta = 0$, consumers are blind with respect to government activity. Government bonds are then completely perceived as a part of permanent income. Consumers have an infinite horizon but disregard (future) government activity. In the more realistic case that $0<\theta<1$, the consumers only partially incorporate future taxes and government expenditures and perceive bonds partially as wealth.

When $\rho>0$ and $\theta>0$ there is an additional wealth effect entering (5), captured by the term $\sum_{j=0}^{\infty}(1+r+\rho)^{-(j+1)}(1+r)^{-1}\theta\rho E_t b_{t+1+j}$. Expected future government debt is, to some extent, considered wealth, even if $\theta = 1$. The reason is that, for a given path of government expenditures, current debt will imply either future taxes or more future debt. Since the consumers have a shorter horizon than the government, future taxes implied by current government debt are not completely taken into account. Rather, there is a chance that government debt will not have to be paid for by future taxes. Future government debt can then be considered (partially) as wealth. The higher the expected future debt path and hence the higher $E_t b_{t+1}$, $E_t b_{t+2}, \ldots$, the higher consumption. If $\rho>0$ and $\theta=0$, on the other hand, consumers have a short horizon and do not look at government activity so that government bonds are completely perceived as wealth.

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[5] On the other hand there may also be combinations of $\lambda\neq 0$ and $\theta<1$ which are compatible with Ricardian Equivalence. To illustrate this with an extreme example, assume that taxes were always equal to government expenditures. Then choosing $\theta=1$ and $\lambda=0$ in (3) would be equivalent to choosing $\theta = 0.5$ and $\lambda=-1$.

To remove the unobservable components from (5) we use the following first-order difference equations (see Hayashi 1982),

$$(6)\quad \sum_{j=0}^{\infty}(1+r+\rho)^{-j}E_t y_{t+j}=(1+r+\rho)\left[\sum_{j=0}^{\infty}(1+r+\rho)^{-j}E_{t-1}y_{t-1+j}-y_{t-1}\right]+e_{yt}$$

$$(7)\quad \sum_{j=0}^{\infty}(1+r+\rho)^{-j}E_t g_{t+j}=(1+r+\rho)\left[\sum_{j=0}^{\infty}(1+r+\rho)^{-j}E_{t-1}g_{t-1+j}-g_{t-1}\right]+e_{gt}$$

$$(8)\quad \sum_{j=0}^{\infty}(1+r+\rho)^{-j}E_t b_{t+1+j}=(1+r+\rho)\left[\sum_{j=0}^{\infty}(1+r+\rho)^{-j}E_{t-1}b_{t+j}-b_t\right]+e_{bt}$$

where $e_{yt}=\sum_{j=0}^{\infty}(1+r+\rho)^{-j+1}\left(E_t y_{t+j-1}-E_{t-1}y_{t+j-1}\right)$, $e_{gt}=\sum_{j=0}^{\infty}(1+r+\rho)^{-j+1}\left(E_t g_{t+j-1}-E_{t-1}g_{t+j-1}\right)$

and $e_{bt}=\sum_{j=0}^{\infty}(1+r+\rho)^{-j+1}\left(E_t b_{t+j}-E_{t-1}b_{t+j}\right)$. These errors denote the revisions of expectations made by the consumers as they move from t-1 to t. They are uncorrelated with variables known at time t-1, but may be correlated with variables dated period t and contemporaneously correlated with each other. For an explanatory note on these difference equations we refer to Appendix A2. Equations (6), (7) and (8) together with (2) are used in (5) to obtain a consumption function that can be estimated (see Appendix A2),

$$(9)\quad \begin{aligned} c_t &= (1+r+\rho)c_{t-1}+\theta\beta(1-\lambda)(1+r+\rho)g_{t-1}-\lambda(g_t-(1+r+\rho)g_{t-1})-\beta\theta\rho(1+r)^{-1}b_t \\ &+\beta(1-\theta)(b_t-(1+r+\rho)b_{t-1})+\beta(w_t-(1+r+\rho)w_{t-1})-\beta(1+r+\rho)y_{t-1}+u_t \end{aligned}$$

where $u_t=v_t-(1+r+\rho)v_{t-1}+\beta e_{yt}-\theta\beta(1-\lambda)e_{gt}+\beta\theta\rho(1+r+\rho)^{-1}(1+r)^{-1}e_{bt}$.

3. Methodology.

In this section we discuss the data to be used in the estimation of the consumption function derived in the last section. We also discuss the estimation method. Special attention goes to the methodological problems that arise because the consumption function is expressed in levels whereas the included variables may be non-stationary. Estimation results are then presented in the next section.

3.1 Data considerations.

This section describes the data that we use to estimate equation (9). For private consumption (c_t) we follow among others Aschauer (1985), Evans (1988) and Graham and Himarios (1991, 1996). We use consumer expenditures on non-durable consumer goods and services [6]. For g_t we use the sum of government consumption (wage and non-wage consumption) and government investment. We proxy private financial wealth w_t at the beginning of the year by the capital stock of the business sector (see Perelman and Pestieau 1993) augmented with net foreign assets. For b_t we use *net* government debt since rational consumers will most likely incorporate the government's *net* claims on the private sector. For y_t we use pre-tax labour income augmented with transfers[7]. To calculate pre-tax labour income, government wage consumption is added to the product of the wage rate and employment in the business sector. For transfers we use social security benefits

[6] The most appropriate concept for consumption however is expenditures on non-durables and services augmented with a service flow from consumer durables. The latter is difficult if not impossible to calculate for Belgium for reasons of data availability.
[7] If we do not include transfers, the conclusions are the same.

paid by the government. We do not take into account labour income of the self-employed, due to the difficulties in imputing labour earnings to the self-employed. Equation (9) is estimated using annual data for Belgium from 1970 to 1997 [8]. Data and data sources are described in table 1.

Table 1. Data and data sources.

c_t	Real per capita expenditures on non-durable goods and services. Calculated from several volumes from 'Statistisch Jaarboek van België' (NIS) for period 1970-1979. Calculated from 'Nationale Rekeningen' (INR, 1998) for period 1980-1997. Calculated as the difference between total conumption and expenditures on durables (including cars). Deflated by implicit price deflator for non-durables and services (1990=100).
g_t	Real per capita government consumption plus real per capita government investment. Taken from OECD Economic Outlook (CD ROM 1999 Vol. 1) and reported in real terms with codes CGV for real government consumption and IGV for real government investment (thus deflated by implicit price deflators for government consumption and government investment respectively, 1990=100).
w_t	Real per capita capital stock of the business sector plus real per capita net foreign assets. Capital Stock of the business sector is reported in real terms in OECD Economic Outlook (CD ROM 1999 Vol.1) with code KBV (1990=100) and net foreign assets is from Lane and Milesi-Ferretti (1999, 2001a, 2001b). For more on the construction of net foreign assets we refer to appendix B. This series is deflated by the implicit price deflator for non-durables and services (1990=100).
b_t	Real per capita net government debt. Nominal net government debt is taken from OECD Economic Outlook (CD ROM 1999 Vol. 1) with code GNFL. This series is deflated by the implicit price deflator for non-durables and services (1990=100).
y_t	Real per capita pre-tax labour income. Obtained by adding the product of the wage rate of the business sector and employment in the business sector to social security benefits paid by the government. These variables are from OECD Economic Outlook (CD ROM 1999 Vol.1) and the codes are WR for wage rate business sector, (EE-EG) for employment business sector and SSPG for social security benefits paid by the government. The resulting series is deflated by the implicit price deflator for non-durables and services (1990=100).

Note : per capita measures are obtained after dividing by total population (from OECD Economic Outlook CD ROM 1999 Vol.1 with code POP).

3.2 Methodology.

Since equation (9) is non-linear in its parameters (ρ, λ, θ and β) we use a non-linear estimation method. To take into account the possibility of simultaneity, we estimate (9) by non-linear instrumental variables (NLIV). Simultaneity can for instance occur if news in period t about an increase in government spending in the future (a positive shock in e_{gt} which is part of the error term u_t) coincides with an increase in government spending in period t; thus g_t and u_t could be correlated. Moreover, it is possible that transitory consumption (v_t) is correlated with labour

[8] Although this is a short period for time series analysis, it is certainly no exception in the literature (see e.g. Nicoletti, 1988 ; Evans, 1993 ; Giavazzi and Pagano 1990, 1995 ; Graham and Himarios, 1996) .

income and wealth. To obtain consistent estimates we need to use instruments. The most obvious choice, which also excludes data mining, is to use the lagged values of the variables appearing in (9) as instruments. Since the model implies autocorrelation of the MA(1) form in the error term u_t, the use of first lags (dated period t-1) is not valid. We therefore use a constant and the second lags of the variables c_t, g_t, w_t, b_t and y_t as instruments.

It is common practice to estimate equations of the same type as (9) in levels [9]. However, this approach does not take into account that with non-stationary variables expressed in levels, standard inference procedures are not valid since the estimators do not necessarily have the standard asymptotic distribution. Note that (9) can also be written as,

$$\begin{aligned}\Delta c_t = &-\lambda\Delta g_t + \beta\Delta w_t + (\beta(1-\theta) - \beta\theta\rho(1+r)^{-1})\Delta b_t + [\ (r+\rho)c_{t-1} - \beta(r+\rho)w_{t-1} \\ &- \beta(1+r+\rho)y_{t-1} + (\lambda(r+\rho) + \theta\beta(1-\lambda)(1+r+\rho))g_{t-1} \\ &- \beta(r(1-\theta) + \rho(1+r(1-\theta))(1+r)^{-1})b_{t-1}\ ] + u_t\end{aligned} \tag{9'}$$

Sims, Stock and Watson (1990) argue that if a regression equation with variables written in levels can be reparameterized such that certain parameters can be expressed as coefficients only on I(0) variables, then those coefficients will have the standard asymptotic distribution. Assuming that g_t and w_t, for instance, are I(1) variables, we might conclude from this that β and λ have the asymptotic normal distribution [10]. However, due to the highly non-linear form of (9) and (9'), the

[9] See for instance Giavazzi and Pagano (1995) and Graham and Himarios (1996).

[10] Using augmented Dickey-Fuller unit root tests (not reported), we cannot reject that the levels of our variables contain a unit root. This may be due to the small sample size, which can considerably reduce the power of unit root tests. Note from (9') that assuming c_t, g_t, w_t, b_t and y_t are I(1) variables and assuming the error is stationary, the theoretical model implies that there is a cointegrating relationship between c_t, g_t, w_t, b_t and y_t (i.e the term in square brackets is stationary).

estimators of all coefficients are linked and do all have a non-standard asymptotic distribution. For instance, β and λ also appear as coefficients on I(1) variables. Therefore it is useful to simulate the small sample distributions for all parameters through a bootstrap procedure. The critical values of these bootstrapped distributions can then be used to test hypotheses on the parameters of interest.

Note that since we work with a small sample, the bootstrap procedure that we implement would be useful, even if there were no problems of stationarity. It is often argued that when using (low quality) instrumental variables in a small sample, estimated parameters and their standard errors may be biased and the assumed asymptotic distribution may be a poor approximation to the actual finite sample distribution (see Nelson and Startz 1990).

To calculate small sample distributions we conduct a parametric bootstrap through *residual resampling*. This has the advantage that we do not have to impose a normality assumption on the error term to generate residuals. The error term in (9) has an MA(1) component. In our case (given the complicated structure of the error term) the most obvious way to take into account autocorrelation is through the resampling of blocks of residuals instead of individual residuals (see Maddala and Kim 1998). Our methodology is explained in more details in Appendix C.

4. Results.

In this section we discuss the results of the estimation of the consumption function. Robustness tests are also presented. The section ends with a summary of the results.

4.1 Basic regression results.

The results from estimating (9) with NLIV are presented in table 2. We set r=0.03, r=0.05 and r=0.07. These values are in the range considered by Evans (1993) [11].

We can see from table 2 that the estimates of all the coefficients have correct signs in all the equations. The estimated values for ρ, λ and β also make economic sense. Using the critical values of the standard normal distribution we would conclude that the estimates for β and ρ are significantly different from zero at the 5% level in all equations; λ is never significantly different from zero.

Theoretically the parameter θ must lie in the interval [0,1]. The point estimate for θ is close to 1 when r=0.07, but somewhat larger than 1 when r=0.03 and r=0.05. Note however that the standard errors on this coefficient are much larger than those on the estimates for ρ and β. The parameter θ is thus estimated relatively imprecisely. Using the standard normal distribution we can reject θ=0 (at the 5% level of significance using a two-sided test) in all equations but we can never reject θ=1. We cannot reject that θ takes on a certain number of values between 0 and 1 either. We can thus reject complete but not partial myopia with respect to government activity.

Of course, using the normal distribution for inference may not be the correct approach, so we use bootstrap methods to generate small sample distributions (see appendix C). In table 2 we report the conclusions of testing $\theta=0$ vs. $\theta>0$, $\rho=0$ vs. $\rho>0$, $\beta=0$ vs. $\beta>0$ (one-sided tests) and $\lambda=0$ vs. $\lambda\neq0$ (two-sided tests) using the critical values of the small sample distribution of the *point estimates* of the parameter under consideration. These critical values are reported in appendix C.

[11] Note for instance that the real interest rate on 10 year Belgian government bonds averaged 4.5 % over the period 1970-1997 (see OECD Economic Outlook).

Table 2. Estimation of (9) by Non Linear Instrumental Variables (NLIV) [a], annual data for Belgium (1970-1997).

	(1)	(2)	(3)
	r=0.03	r=0.05	r=0.07
ρ	0.121 * (0.036)	0.117 * (0.043)	0.112 ** (0.050)
λ	-0.430 (0.576)	-0.499 (0.544)	-0.553 (0.513)
θ	1.226 ** (0.506)	1.118 ** (0.484)	1.020 ** (0.475)
β	0.084 ** (0.015)	0.087 ** (0.017)	0.088 ** (0.018)
N obs	26	26	26
R^2adj.	0.990	0.990	0.989
DW	1.885	1.867	1.850
BG(1)	0.772	0.735	0.701

Notes: Newey-West standard errors between brackets (lag truncation = 3). DW denotes the Durbin-Watson test statistic. BG(1) denotes the p-value of the Breusch-Godfrey test for serial correlation of order 1 (the null hypothesis is no autocorrelation). N obs denotes the number of observations used in the estimation (after taking into account the loss of two observations to form instruments). [a] used instruments are a constant and the second lag of consumer expenditures on non-durables and services, government expenditures, private financial wealth, net government debt and pre-tax labour income augmented with transfers. * (**) denotes that the estimate for the coefficient under consideration is significant at the 10% (5%) level using the critical values of the bootstrapped distributions. We use the critical values of the distribution of the *point estimates* under the null (see appendix C). For the parameters ρ, θ and β we use the 10 % (5%) critical values of a one-sided test for ρ=0 against ρ>0, θ=0 against θ>0 and β=0 against β>0. For λ we use the 10% (5%) critical values of a two-sided test for λ=0 against $\lambda \neq 0$. Note that θ=1 is never rejected using bootstrapped distributions (at the 10% level of significance using both a one-sided test θ=1 against θ<1 and a two-sided test θ=1 against $\theta \neq 1$).

The results of using the small sample distribution of the *t-values* instead (also reported in appendix C) do generally point in the same direction, though the levels of significance of the results may differ. We prefer using point estimates since t-values are constructed using (asymptotic) Newey-West standard errors. The latter may be unreliable[12]. The use of small sample distributions for inference largely confirms the results found using the (asymptotic) normal distribution. The significance of the results is generally lower however. The parameter λ is never significantly different from zero. β is significantly larger than zero at the 5% level of significance. ρ is significantly larger than zero at the 10% level of significance in equations (1) and (2) and at the 5% level in equation (3). The parameter θ is significantly larger than zero at the

[12] Our simulations suggest that the asymptotic standard error of a coefficient can sometimes differ considerably from its bootstrapped standard error, especially for the parameters ρ and θ (unreported).

5% level in all equations. We can never reject that θ is taken from the $\theta=1$ distribution (both with a one-sided test $\theta=1$ vs. $\theta<1$ and a two-sided test $\theta=1$ vs. $\theta\neq1$, at the 10% level of significance). Note that using the 10% critical values of the *t-ratios* we cannot reject that $\theta=0$. Using t-ratios we can reject $\theta=0$ at the 13% level in equation (1) and at the 15% level in equations (2) and (3) however (unreported).

Since we cannot reject both $\theta=1$ and $\lambda=0$, it is interesting to re-estimate equation (9) imposing these restrictions to know how the estimates of the remaining parameters are affected. The results of the restricted estimation of (9) with r=0.05 are presented in table 3.

Table 3. Restricted estimation of (9) with r=0.05 using NLIV [a], annual data for Belgium (1970-1997).

	(1)	(2)	(3)
	$\theta=1$	$\lambda=0$	$\theta=1$ and $\lambda=0$
ρ	0.124 ** (0.035)	0.116 * (0.027)	0.118 ** (0.026)
λ	-0.516 (0.383)	0	0
θ	1	1.257 * (0.350)	1
β	0.084 ** (0.007)	0.094 ** (0.010)	0.089 ** (0.006)
N obs	26	26	26
R^2adj.	0.990	0.990	0.991
DW	1.877	1.683	1.667
BG(1)	0.753	0.399	0.419

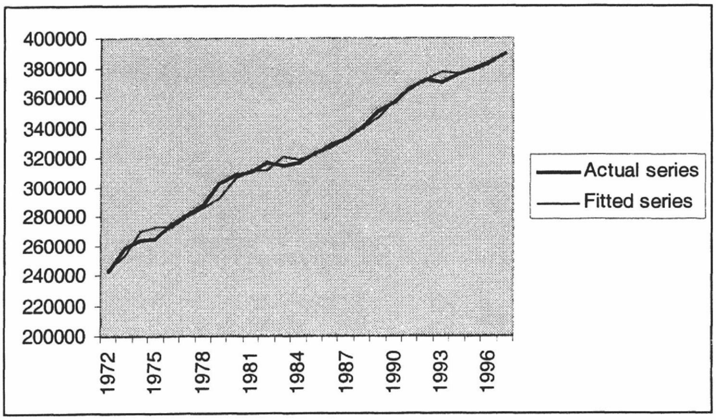
Notes: Newey-West standard errors between brackets (lag truncation = 3). DW denotes the Durbin-Watson test statistic. BG(1) denotes the p-value of the Breusch-Godfrey test for serial correlation of order 1 (the null hypothesis is no autocorrelation). N obs denotes the number of observations used in the estimation (after taking into account the loss of two observations to form instruments). [a] used instruments are a constant and the second lag of consumer expenditures on non-durables and services, government expenditures, private financial wealth, net government debt and pre-tax labour income augmented with transfers. * (**) denotes that the estimate for the coefficient under consideration is significant at the 10% (5%) level using the critical values of the bootstrapped distributions. We use the critical values of the distribution of the *point estimates* under the null. For the parameters ρ, β and θ (the latter only in equation (2)) we use the 10 % (5%) critical values of a one-sided test for $\rho=0$ against $\rho>0$, $\theta=0$ against $\theta>0$ and $\beta=0$ against $\beta>0$. For λ (in equation (1)) we use the 10% (5%) critical values of a two-sided test for $\lambda=0$ against $\lambda\neq0$. Note that in equation (2) $\theta=1$ is not rejected using bootstrapped distributions (at the 10% level of significance using both a one-sided test $\theta=1$ against $\theta<1$ and a two-sided test $\theta=1$ against $\theta\neq1$).

As can be seen, comparing equation (2) in table 2 to its restricted counterparts in table 3, the point estimates of the remaining parameters do generally not change very much (note that this conclusion remains valid if we impose other values for r in the range r=0.03 to r=0.07). An exception is the point estimate of θ under $\lambda=0$, which rises considerably. Using bootstrapped distributions it is only significantly larger than zero at the 10% level this time. The discount rate ρ is now significantly larger than zero at the 5% level if $\theta=1$ or if $\theta=1$ and $\lambda=0$.

4.2 Robustness tests.

A first additional test to see if the model provides a good fit is presented in figure 1 where we show a graph of actual and fitted real per capita expenditures on non-durables and services obtained from estimating equation (9) with r=0.05. The graph suggests that the model gives quite a good approximation to reality. The fit remains good for other values of r between 0.03 and 0.07.

Figure 1. Real per capita consumer expenditures on non-durables and services: actual and fitted (in Belgian francs of 1990)



Note: the fitted series is calculated using the estimation results reported in table 2 (eq. (2) with r=0.05).

We can also check whether our estimation results are sensitive to the use of 1) a different instrument set (with a second *and* third lag of all variables), 2) different variables (government consumption instead of government consumption and investment[13]) and 3) a different sample period (1970-1995 instead of 1970-1997 [14]). The results are presented in table 4.

Table 4. Estimation of (9) with r=0.05 using a different instrument set (1), a different variable for g_t (2) and a different sample period (3).

	(1)	(2)	(3)
ρ	0.116 *	0.131 **	0.124 *
	(0.040)	(0.035)	(0.062)
λ	-0.230	-0.766	-0.443
	(0.538)	(0.621)	(0.527)
θ	1.095 **	1.020	1.048 **
	(0.525)	(0.547)	(0.591)
β	0.088 **	0.082 **	0.087 **
	(0.018)	(0.023)	(0.017)
N obs	25	26	24
R^2adj.	0.988	0.990	0.987
DW	1.729	1.808	1.857
BG(1)	0.601	0.623	0.740

Notes: Newey-West standard errors between brackets (lag truncation = 3). DW denotes the Durbin-Watson test statistic. BG(1) denotes the p-value of the Breusch-Godfrey test for serial correlation of order 1 (the null hypothesis is no autocorrelation). N obs denotes the number of observations used in the estimation (after taking into account the loss of observations to form instruments). In eq. (1) the instrument set is the same as in table 2 except that an additional third lag of all variables is added. Eq. (2) differs from the results of table 2 because we use government consumption instead of government consumption and investment. In eq (3) we use a smaller sample period (1970-1995) than in table 2.* (**) denotes that the estimate for the coefficient under consideration is significant at the 10% (5%) level using the critical values of the bootstrapped distributions. We use the critical values of the distribution of the *point estimates* under the null. For the parameters ρ, θ and β we use the 10 % (5%) critical values of a one-sided test for ρ=0 against ρ>0, θ=0 against θ>0 and β=0 against β>0. For λ we use the 10% (5%) critical values of a two-sided test for λ=0 against λ≠0. Note that θ=1 is never rejected using bootstrapped distributions (at the 10% level of significance using both a one-sided test θ=1 against θ<1 and a two-sided test θ=1 against θ≠1).

[13] Even though the Ricardian Equivalence debate suggests to use government consumption and investment (i.e. the government budget constraint), the Edgeworth concept (which allows for a role of government expenditures in the utility function) is frequently tested using only government consumption.

[14] Using a reduced sample size is useful to make sure that the results are not dependent on one or more observations. We use the sample period 1970-1995 because during this period the real per capita net government debt increased almost uninterruptedly while it fell in both 1996 and 1997. Note that using other reduced samples (e.g 1971-1997; 1970-1996;...) does not affect our conclusions either.

As can be seen from table 4 the results are fairly robust. If we compare the results in table 4 with the result found in table 2 (eq. (2)) we can see that the point estimates do not change much. The significance of the results is only altered in eq.(2) of table 4 where we cannot reject that θ is zero. Note that these conclusions generally also hold if r=0.03 or if r=0.07.

4.3 Summary of the results.

Summarizing, our results for λ are consistent with the hypothesis that private consumption decisions are independent of expected government expenditures, conditional on expected taxes. The value for θ is in general significantly different from zero but not significantly different from one, suggesting that Belgian consumers are not myopic with respect to government activity. Graham and Himarios (1991) on the other hand find, in their framework, that the parameter of myopia is small and not significantly different from zero in the US.

The point estimate of the mark-up ρ seems to be larger than zero which indicates that the time horizon of Belgian consumers might be shorter than that of the government. They thus discount future income (and future taxes, since they are not myopic with respect to government activity) at a rate higher than the real rate of return. Alternatively, we can argue that Belgian consumers are prudent and save out of precaution. The magnitude of the estimates for ρ and for the discount rate $r+\rho$ is consistent with those of previous studies. Hayashi (1982) finds values of the discount rate of around 13%. Graham and Himarios (1991) find estimates between 12% and 21%. Graham and Himarios (1996) even find estimates of around 35%. Friedman (1957) suggested that the discount rate might be as high as one-third (see also Muellbauer and Lattimore 1995).

5. Conclusions.

Our conclusions support the hypothesis that Belgian consumers are not myopic with respect to (future) government activity. In addition, we cannot reject that Belgian consumers have a relatively high overall discount rate. This may be because they have a general planning horizon that is shorter than the government's horizon. Alternatively, we can interpret this as an indication that Belgian consumers have a precautionary savings motive. These results imply that the *complete* incorporation of the future tax implications of current debt, and thus Ricardian Equivalence, is rejected. Our results nevertheless suggest a *partial* incorporation by the consumers of future tax implications of debt. Given these results fiscal policy will be effective but less effective than if the consumers were more (or, as an extreme case, completely) myopic with respect to (future) government activity. The result is in line with previous findings in the literature which suggest that the Ricardian Equivalence hypothesis may hold only partially (see e.g. Doménech et. al.). Finally, we cannot reject that private consumption and government expenditures are Edgeworth independent.

The result that Belgian consumers may be aware of (future) government activity seems plausible, given the problematic debt situation in Belgium, especially in the 1980s and 1990s. It is in line with the view expressed in the literature that consumers in countries with unsustainable debt paths tend to be more aware of the future tax implications of this debt (see Nicoletti 1988, 1992; Sutherland 1997). Unlike Nicoletti (1988) we do not find complete incorporation of future taxes, since our results suggest that $\rho>0$. The discount rate $r+\rho$ is an overall rate that reflects many things that are not necessarily related to government issues (risk aversion, variance of future pre-tax labour income). It is likely that the parameter θ, on the other hand, is more closely linked to debt issues.

Our conclusions depend strongly on the estimation of the parameter capturing myopia with respect to government activity. The estimation of this parameter and the estimation of the consumption function as a whole are not without problems. They are complicated because of stationarity concerns, endogeneity and sample size. The use of small sample distributions can to a certain extent alleviate these problems. In many cases the results obtained using these distributions have lower statistical significance. The weight given to our results must take into account these empirical limitations.

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Appendix A1: derivation of equation (5).

To derive (5) we first write (4) in the main text for period t+j instead of t. We then multiply both sides of the equation by $\theta(1+r+\rho)^{-j}$, and take the sum from j=1 to j=∞ of all the terms. Taking the expectations of both sides at time t and rearranging, we then obtain,

$$\text{(A1.1)}\quad \begin{aligned} &\theta\sum_{j=1}^{\infty}(1+r+\rho)^{-j}E_t t_{t+j} = \theta\sum_{j=1}^{\infty}(1+r+\rho)^{-j}E_t g_{t+j} + \theta\sum_{j=1}^{\infty}(1+r+\rho)^{-j}E_t b_{t+j} \\ &-\theta(1+r)^{-1}\sum_{j=1}^{\infty}(1+r+\rho)^{-j}E_t b_{t+1+j} \end{aligned}$$

Adding and subtracting the term $\theta(1+r)^{-1}E_t b_{t+1}$ to the RHS of this equation yields,

$$\begin{aligned} &\theta\sum_{j=1}^{\infty}(1+r+\rho)^{-j}E_t g_{t+j} + \theta\sum_{j=1}^{\infty}(1+r+\rho)^{-j}E_t b_{t+j} - \theta(1+r)^{-1}\sum_{j=1}^{\infty}(1+r+\rho)^{-j}E_t b_{t+1+j} \\ &+\theta(1+r)^{-1}E_t b_{t+1} - \theta(1+r)^{-1}E_t b_{t+1} \end{aligned}$$

From (4) we have that $\theta(1+r)^{-1}E_t b_{t+1} = \theta(g_t - t_t + b_t)$ such that we get,

$$\text{(A1.2)} \quad \begin{aligned} &\theta\sum_{j=1}^{\infty}(1+r+\rho)^{-j}E_t t_{t+j} = \theta\sum_{j=1}^{\infty}(1+r+\rho)^{-j}E_t g_{t+j} + \theta\sum_{j=1}^{\infty}(1+r+\rho)^{-j}E_t b_{t+j} \\ &-\theta(1+r)^{-1}\sum_{j=1}^{\infty}(1+r+\rho)^{-j}E_t b_{t+1+j} + \theta(g_t - t_t + b_t) - \theta(1+r)^{-1}E_t b_{t+1} \end{aligned}$$

The newly introduced terms $\theta(g_t - t_t)$ can be brought into the summation signs to get,

$$\text{(A1.3)} \quad \begin{aligned} &\theta\sum_{j=0}^{\infty}(1+r+\rho)^{-j}E_t t_{t+j} = \theta\sum_{j=0}^{\infty}(1+r+\rho)^{-j}E_t g_{t+j} + \theta b_t \\ &+\theta\sum_{j=1}^{\infty}(1+r+\rho)^{-j}E_t b_{t+j} - \theta(1+r)^{-1}\sum_{j=1}^{\infty}(1+r+\rho)^{-j}E_t b_{t+1+j} - \theta(1+r)^{-1}E_t b_{t+1} \end{aligned}$$

The expression on the last line in (A1.3) equals

$$+\theta(1+r+\rho)^{-1}\sum_{j=1}^{\infty}(1+r+\rho)^{-j+1}E_t b_{t+j} - \theta(1+r)^{-1}\sum_{j=0}^{\infty}(1+r+\rho)^{-j}E_t b_{t+1+j}$$

and can be written as $\theta\left[(1+r+\rho)^{-1} - (1+r)^{-1}\right]\sum_{j=0}^{\infty}(1+r+\rho)^{-j}E_t b_{t+1+j}$. Using the fact that

$\left[(1+r+\rho)^{-1} - (1+r)^{-1}\right] = \dfrac{(1+r) - (1+r+\rho)}{(1+r+\rho)(1+r)} = -\dfrac{\rho}{(1+r+\rho)(1+r)}$ we can then write (A1.3) as,

$$\text{(A1.4)} \quad \begin{aligned} &\theta\sum_{j=0}^{\infty}(1+r+\rho)^{-j}E_t t_{t+j} = \theta\sum_{j=0}^{\infty}(1+r+\rho)^{-j}E_t g_{t+j} + \theta b_t \\ &-\theta\rho(1+r+\rho)^{-1}(1+r)^{-1}\sum_{j=0}^{\infty}(1+r+\rho)^{-j}E_t b_{t+1+j} \end{aligned}$$

Substituting (A1.4) in (3), we obtain (5) in the main text.

Appendix A2: a note on the difference equations (6), (7) and (8) - derivation of equation (9).

A note on the difference equations (6), (7) and (8).

In this appendix we show how we obtain equations (6), (7) and (8). Suppose we have a variable x_t and a discount rate m. Then we can write,

$$\text{(A2.1)}\quad \sum_{j=0}^{\infty}(1+m)^{-j}x_{t+j} = (1+m)\left(\sum_{j=0}^{\infty}(1+m)^{-j}x_{t-1+j} - x_{t-1}\right)$$

First, we take expectations at time t of both sides of (A2.1). Then, we add *and* subtract the term $(1+m)\sum_{j=0}^{\infty}(1+m)^{-j}E_{t-1}x_{t-1+j}$ at the RHS of (A2.1) to obtain,

$$\text{(A2.2)}\quad \sum_{j=0}^{\infty}(1+m)^{-j}E_t x_{t+j} = (1+m)\sum_{j=0}^{\infty}(1+m)^{-j}E_{t-1}x_{t-1+j} - (1+m)x_{t-1} + e_{xt}$$

where $e_{xt} = (1+m)\left(\sum_{j=0}^{\infty}(1+m)^{-j}E_t x_{t-1+j} - \sum_{j=0}^{\infty}(1+m)^{-j}E_{t-1}x_{t-1+j}\right)$

Now replace x_t with y_t, g_t and b_{t+1} and m with r+ρ in (A2.2) to obtain equations (6), (7) and (8) in the main text.

Derivation of equation (9).

Equation (5) can also be written as,

$$\text{(A2.3)}\quad \begin{aligned} c_t^* &= \beta w_t + \beta(1-\theta)b_t + \beta\sum_{j=0}^{\infty}(1+r+\rho)^{-j}E_t y_{t+j} - \theta\beta(1-\lambda)\sum_{j=0}^{\infty}(1+r+\rho)^{-j}E_t g_{t+j} \\ &+ \beta\theta\rho(1+r+\rho)^{-1}(1+r)^{-1}\sum_{j=0}^{\infty}(1+r+\rho)^{-j}E_t b_{t+1+j} + v_t \end{aligned}$$

We lag equation (A2.3) one period so that we can derive,

(A2.4)

$$\beta\sum_{j=0}^{\infty}(1+r+\rho)^{-j}E_{t-1}y_{t-1+j} = c_{t-1}^* - \beta w_{t-1} - \beta(1-\theta)b_{t-1}$$

$$+\theta\beta(1-\lambda)\sum_{j=0}^{\infty}(1+r+\rho)^{-j}E_{t-1}g_{t-1+j} - \beta\theta\rho(1+r+\rho)^{-1}(1+r)^{-1}\sum_{j=0}^{\infty}(1+r+\rho)^{-j}E_{t-1}b_{t+j} - v_{t-1}$$

Using (6), (7) and (8) into (A2.3), we can write,

$$\text{(A2.5)}\quad \begin{aligned} c_t^* &= \beta\left[(1+r+\rho)\left(\sum_{j=0}^{\infty}(1+r+\rho)^{-j}E_{t-1}y_{t-1+j} - y_{t-1}\right) + e_{yt}\right] \\ &-\theta\beta(1-\lambda)\left[(1+r+\rho)\left(\sum_{j=0}^{\infty}(1+r+\rho)^{-j}E_{t-1}g_{t-1+j} - g_{t-1}\right) + e_{gt}\right] \\ &+ \beta\theta\rho(1+r+\rho)^{-1}(1+r)^{-1}\left[(1+r+\rho)\left(\sum_{j=0}^{\infty}(1+r+\rho)^{-j}E_{t-1}b_{t+j} - b_t\right) + e_{bt}\right] \\ &+ \beta w_t + \beta(1-\theta)b_t + v_t \end{aligned}$$

Plugging (A2.4) into (A2.5) and using (2) makes it possible to get rid of the unobservable terms. After some rearrangements this leads to equation (9) in the main text.

Appendix B : A note on the external wealth data.

Lane and Milesi-Ferretti (1999, 2001a, 2001b) proxy net foreign assets by cumulating current account surpluses with adjustments. The most recent data can be found on Philip Lane's data page, internet site URL: http://econserv2.bess.tcd.ie/plane/data.html. The net foreign assets series is proxied by the ACUMCA (adjusted cumulative current account) series. This series is reported for 1970-1997 in USD and for Belgium and Luxembourg together. It is not reliable, however, for the nineties (see notes and definitions accompanying the data table on the website). Gian Maria Milesi-Ferretti has suggested to use an alternative measure for the adjusted cumulative current account, namely a series that does not adjust for valuation changes in portfolio equity (the series ACUMCA-CEQAR+CEQA+CEQLR-CEQL - for definitions we refer to the data set on the website). Unlike the ACUMCA series, this series does not show the implausible negative values in the nineties.

To use this series in our regressions we express it in BEF using USD-BEF purchasing power parities from OECD Economic Outlook. Since current account data and therefore this net foreign assets series is only available for Belgium *and* Luxembourg together, we can only approximate it imperfectly for Belgium. We do this by multiplying our net foreign assets series that holds for Belgium-Luxembourg by the ratio of GDP for Belgium to GDP for Belgium and Luxembourg (the average of this ratio over the period 1970-1997 is 1.05)

Appendix C : Generating bootstrapped distributions.

We use the moving block bootstrap where *overlapping* blocks of the residuals of the estimation of (9) are resampled with replacement. We refer to Maddala and Kim (1998) for the advantages of this procedure compared to the non-overlapping block bootstrap procedure. We consider blocks with two observations, thus block length le = 2 (see Maddala and Kim 1998). Note that our conclusions remain the same if le=3 or if le=1 (the latter case is a standard bootstrap).

To carry out the bootstrap, for example to test whether $\theta=0$ in the three cases considered in table 2, we proceed as follows (for le=2):

1) We estimate (9) under the restriction $\theta = 0$ to obtain ρ , $\hat{\lambda}$, β and the residuals u_{ot}. We divide the obtained series u_{ot} , containing 26 usable observations (28 from the sample size 1970-1997 minus 2 necessary to form instruments), into blocks of length 2 (the first block contains observations 1 and 2, the second 2 and 3, and so on). We then have 25 blocks to choose from.

2) We draw a sample of size 26 by randomly selecting 13 of these blocks *with replacement*. We then obtain a pseudo-time series of residuals e_{ot} by laying the blocks end-to-end in the order sampled.

3) Using the actual time series for g_t , n_t , b_t , y_t , the first data point of c_t, the error series e_{ot} and the parameter values $\rho = \rho$, $\lambda = \hat{\lambda}$, $\beta = \beta$, $\theta = 0$, we generate a pseudo data series $c_t^{'}$ through (9).

We then use this series $c_t^{'}$ together with g_t , n_t , b_t , y_t to estimate (9) with NLIV. We obtain an estimate for the parameter θ and its standard error. We calculate the t-ratio as well.

4) We repeat steps 2 and 3 a large number of times (we use 10 000 replications in each case). Then we order the point estimates for θ and the related t-ratios, which results in the small sample distribution of θ and the distribution of the t-ratio under θ=0.

We can test whether the point estimate for θ reported in table 2 is taken from this distribution 1) by looking at the distribution of the point estimates (comparing the point estimate reported in table 2 with critical values from the distribution of the point estimates) or 2) by looking at the distributions of the t-ratios (comparing the t-ratio calculated from table 2 with the critical values from the distribution of the t-ratio). The empirical distribution of the t-values thus replaces the standard normal distribution for inference. In table C1 we report the medians and some other percentiles of the bootstrapped distributions that can be used as critical values (both for the distributions of the point estimates - which are used in table 2 – and for the distributions of the t-values). We report these for all parameters in equations (1), (2) and (3) in table 2.

The results in table 3 are obtained in a similar way (also with le=2). For instance to test the hypothesis θ=0 in equation (2) in table 3 where r=0.05 and λ=0 we calculate the small sample distribution of θ and its t-ratio under λ=0. This means that in steps 1-4 the estimated value for λ is replaced by its restricted value 0. The results in table 4 are also with le=2. For instance, in equation 3 of table 4, the use of a reduced sample size (1970-1995) implies that u_{ot} contains only 24 observations. We then randomly draw 12 blocks of length 2 to form a pseudo series of residuals. We do not report the bootstrapped distributions of point estimates and t-ratios used in the tables 3 and 4 but they are available upon request.

Table C1. Bootstrap distributions of point estimates and t-ratios (used for inference in table 2).

	Equation	(1)		(2)		(3)	
		r=0.03		r=0.05		r=0.07	
Distributions	*Percentiles*	*Point estimates*	*t-ratios*	*Point estimates*	*t-ratios*	*Point estimates*	*t-ratios*
Distribution $\rho=0$ [a]	*50*	0.012	0.240	-0.016	-0.377	-0.019	-0.527
	90	0.114	2.568	0.084	1.700	0.061	1.133
	95	0.142	3.391	0.118	2.546	0.102	1.707
Distribution $\lambda=0$ [b]	*5*	-1.002	-2.658	-0.953	-2.494	-0.928	-2.411
	10	-0.817	-2.121	-0.775	-1.990	-0.738	-1.952
	50	-0.164	-0.404	-0.110	-0.267	-0.077	-0.196
	90	0.579	1.137	0.674	1.478	0.714	1.678
	95	0.873	1.522	0.995	2.074	1.029	2.345
Distribution $\theta=0$ [a]	*50*	-0.009	-0.019	-0.008	-0.015	-0.022	-0.042
	90	0.758	2.579	0.814	2.802	0.842	2.828
	95	0.933	3.498	0.992	3.774	1.014	3.887
Distribution $\theta=1$ [c]	*5*	-0.353	-1.895	-0.314	-1.907	-0.277	-1.887
	10	0.046	-1.577	0.076	-1.577	0.131	-1.554
	50	0.958	-0.145	0.969	-0.108	1.005	-0.054
	90	1.475	1.878	1.472	1.992	1.496	2.203
	95	1.654	2.420	1.680	2.552	1.673	2.838
Distribution $\beta=0$ [a]	*50*	-0.037	-1.299	-0.036	-1.268	-0.038	-1.280
	90	0.037	2.164	0.038	2.081	0.036	1.798
	95	0.060	4.077	0.064	3.898	0.064	3.477

Notes : The columns labelled 'point estimates' and 't-ratios' contain percentiles of the bootstrapped distributions of the point estimates and t-ratios of the distribution under consideration. [a] The 5% and 10% critical values for a one-sided test of $\rho=0$ against $\rho>0$, $\theta=0$ against $\theta>0$ and $\beta=0$ against $\beta>0$ are given by the percentiles 95 and 90 of the distributions. [b] Both 10% critical values for a two-sided test $\lambda=0$ against $\lambda\neq0$ are given by the 5 and 95 percentiles. [c] The 5% and 10% critical values for a one-sided test of $\theta=1$ against $\theta<1$ are given by the percentiles 5 and 10 of the distribution. Both 10% critical values for a two-sided test $\theta=1$ against $\theta\neq1$ are given by the 5 and 95 percentiles.

The coefficient of relative risk aversion: a Monte Carlo study investigating small sample estimator problems.

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Abstract.

When estimating consumer Euler equations using the Generalized Method of Moments approach, negative estimates for the coefficient of relative risk aversion are frequently found. In this paper we suggest that this anomaly may be due to a small sample bias. Based on Monte Carlo simulations we show that the estimate of the coefficient of relative risk aversion tends to have a negative bias. Our results suggest that the other estimates in the equation are not (as seriously) biased. Inference is problematic however, since the standard errors of the estimates tend to be underestimated. We illustrate these points through the estimation of an Euler equation using annual data for Belgium.

JEL classifications: C15; D91.

Keywords: Private consumption; Euler equation; risk aversion; Generalized method of moments; Monte Carlo simulation.

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1. Introduction.

Finding a negative or insignificant estimate for the coefficient of relative risk aversion (RRA) or its inverse, the elasticity of intertemporal substitution (IES), is a frequently encountered anomaly in the estimation of consumer Euler equations. Hall (1988) finds insignificant values for the coefficient of RRA/ IES, while Hansen and Singleton (1996) find negative point estimates. Holman (1998) uses annual data to estimate Euler equations in a money in the utility function framework and finds negative estimates for the coefficient of RRA/ IES.

Some authors attribute this anomaly to a flawed specification of the model. Ogaki and Reinhart (1998) for instance show that the exclusion of durable good services and the user cost of consumer durables from the model could be responsible for the bias. Other authors emphasize that the link between the coefficient of RRA and the IES can be broken by the use of a non-expected utility framework or by assuming time-nonseparable preferences (Constantinides, 1990).

In this paper we suggest that this anomaly can also arise as the result of estimation problems. Mao (1990) for instance attributes biases in GMM estimates to small sample problems and measurement errors in consumption data. The small sample properties of GMM estimators in consumption based models have also been studied by Tauchen (1986), Kocherlakota (1990b) and Hansen, Heaton and Yaron (1996) and tend to confirm Mao's findings. New in this paper is that we also look at the small sample behaviour of other estimated parameters in the Euler equation. Some researchers tend to take the poorly estimated coefficient of RRA / IES for granted, and nevertheless draw conclusions from the other parameters in the model (see for instance Holman, 1998). They thus assume that only the coefficient of RRA/ IES is poorly estimated. We construct an Euler equation which allows for the estimation of the coefficient of RRA, but also of a parameter reflecting a potential Edgeworth relationship between private and public consumption (see Aschauer 1985, Karras 1994, Ni 1995) [1] . If the coefficient of RRA/ IES is affected by a small sample bias, the question is then whether both parameters are poorly estimated or whether, on the contrary, the estimation of and inference on the Edgeworth parameter is more reliable than the estimation of the coefficient of RRA/IES.

We also focus on the problem of simultaneously estimating the coefficient of RRA and the subjective rate of time preference.

In section 2 we derive an Euler equation, which allows for a potential Edgeworth relationship between private and public consumption. To capture this relationship, we use a measure that is invariant to increasing linear transformations of the utility function. We estimate this Euler equation by the Generalized Method of Moments (GMM) using annual data for Belgium (1958-1997). In section 3 we conduct Monte Carlo simulations to try to uncover biases and to investigate the small sample distributions of the estimators. Section 4 concludes.

2. Construction and estimation of a consumption Euler equation.

In this section we estimate a consumer Euler equation using annual data for Belgium. We discuss the estimates of the different parameters, including the coefficient of RRA.

2.1 Construction of the Euler equation.

[1] Two goods are Edgeworth substitutes (complements, independent) if increasing the quantity of one good decreases (increases, leaves unaffected) the marginal utility of the other good.

Consider a representative rational consumer who maximizes expected utility by choosing a consumption path over an infinite lifetime,

$$(1)\ \text{Max}\ E_t \sum_{j=t}^{\infty} \alpha^{j-t} U\left[c_j, g_j\right]$$

where α represents a discount factor capturing time-preference ($\alpha > 0$), E_t is the expectations operator conditional on information available at time t, U [ ] is the instantaneous utility function (assumed to be concave), c are private consumer expenditures and g is government consumption. Maximization of (1) subject to a budget constraint (see appendix A1), leads to the following Euler equation describing the agent's optimal consumption plan at time t,

$$(2)\ E_t\left[\alpha(1+r_{t+1})\frac{\partial U/\partial c_{t+1}}{\partial U/\partial c_t} - 1\right] = 0$$

where r_{t+1} is the risk-free interest rate obtained on financial wealth. This equation states that, along the optimal path, the marginal disutility of decreasing consumer expenditures in t by a very small amount will be equal to the expected discounted marginal utility of investing that amount in t and consuming the proceeds in t +1.

We use the term Edgeworth independence to denote the case where the cross - partial derivative of the utility function is zero. If U_{cg} is positive, c and g are Edgeworth complements ; if U_{cg} is negative, c and g are Edgeworth substitutes. Note that if c and g are not Edgeworth independent, a change in g will alter the optimal consumption scheme (through a change in the marginal utility of private consumption). We impose $U(c_t, g_t) = u(c_t^*)$ where c_t^* is effective consumption, which is a function of consumption expenditures (c) and government consumption (g). For u (concave) we use the standard iso-elastic constant coefficient of RRA (CRRA) specification,

$$(3)\quad u(c_t^*) = \frac{1}{\theta}(c_t^{*\theta})$$

where 1-θ is the coefficient of RRA (θ < 1). If c and g are Edgeworth independent this coincides with the reciprocal of the intertemporal elasticity of substitution (IES) of private consumption. We use a Cobb-Douglas specification [2] for effective consumption,

$$(4)\quad c_t^* = c_t^{\beta} g_t^{1-\beta}$$

where the standard assumptions of utility functions are fulfilled if 0<β<1. It then follows that $U(c_t, g_t) = \frac{1}{\theta}\left[c_t^{\beta\theta} g_t^{(1-\beta)\theta}\right]$. From $U_{cg} = \beta(1-\beta)\theta c_t^{\theta\beta-1} g_t^{\theta(1-\beta)-1}$ we can say that if β is equal to 0 or 1 or if θ is equal to 0, c_t and g_t will be independent. Given that 0<β<1, c_t and g_t will be substitutes if θ <0 and c_t and g_t will be complements if θ >0. Notice that β = 1 does not preclude that government consumption enters the utility function. We could for instance set $U(c_t, g_t) = u(c_t^*) + v(g_t)$. This would not change the first order condition.

Since an expected utility function is unique up to an increasing linear transformation, the cross second derivative of a new utility function V defined as an increasing linear transformation of U, will have the same sign as the cross second derivative of U. The

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[2] The choice of a tractable functional form for U(c,g) is of major importance. The standard assumptions that marginal utilities be positive and decreasing must be satisfied. Empirically, the choice must yield Euler equations that are functions of stationary variables. A Cobb-Douglas specification for effective consumption is for instance used by Campbell and Mankiw (1990). Ni (1995) shows that effective consumption is better approximated by a Cobb-Douglas specification than by a linear specification, used for instance by Aschauer (1985) or Karras (1994). We also choose this specification because it is in multiplicative form, which makes the linearizing in section 3 easier.

magnitude of U_{cg} however will have no meaning since it is not invariant to increasing linear transformations of U. We could therefore use a different measure, $-(U_{cg}/U_{cc})$ instead of U_{cg} (with $U_{cc}<0$), which is invariant to positive linear transformations of U and can be derived from the first order condition (2) under certain assumptions [3]. From $-\frac{U_{cg}}{U_{cc}}=\frac{(1-\beta)\theta}{1-\theta\beta}\frac{c_t}{g_t}$, where $1-\theta\beta>0$ for $\theta<1$ and $0<\beta<1$, we can say that for a given value of $\theta \neq 0$, the relationship between c_t and g_t will be weaker the more β approaches 1.

Using (3) and (4) in (2), we can write,

$$(5)\quad E_t\left[\alpha(1+r_{t+1})\left(\frac{c_{t+1}}{c_t}\right)^{\theta\beta-1}\left(\frac{g_{t+1}}{g_t}\right)^{\theta(1-\beta)}-1\right]=0$$

2.2 Estimation of the Euler equation.

The Euler equation (5) is a conditional moment condition of which the parameters can be estimated directly through Hansen's (1982) Generalized Method of Moments (GMM) procedure. To calculate consistent standard errors in the presence of serial correlation and heteroskedasticity in the error term, the estimations are done with Newey and West's (1987) heteroskedasticity and autocorrelation consistent covariance matrix. In a first step we estimate (5) using ordinary instrumental variables. Then we use the first step residuals to weigh the instruments, so that the second step estimation is more efficient. We repeat this proces until

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[3] In steady state and assuming the real rate of return equals the subjective rate of time preference, (2) implies that $U_c(c,g)$ will be constant, i.e the total differential $dU_c(c,g)=U_{cc}dc+U_{cg}dg=0$, so that $dc/dg=-U_{cg}/U_{cc}$.

the estimates converge. It is often argued that this iterative GMM proces is more efficient in small samples than a simple 2-step procedure.

To test the overidentifying restrictions, we use Hansen's (1982) test for overidentifying restrictions. If the population moment conditions are correct, this statistic is distributed chi-squared with degrees of freedom equal to the number of instruments minus the number of parameters to be estimated. It should be emphasized that for the GMM estimator to be efficient, consistent and asymptotically normal, the variables in the Euler equations and the instruments should be stationary. Unit root tests on the variables used confirm that this is indeed the case (see appendix B). As far as the choice of instruments is concerned, Fuhrer, Moore and Schuh (1995) use Monte Carlo simulations to show that the GMM estimator is often biased, statistically insignificant and economically implausible if low-quality instruments are used. A solution to this problem is to use lags of the variables that appear in the Euler equation as instruments.

Moreover, because of potential time aggregation problems (i.e. the frequency with which people make consumption expenditures does not coincide with the frequency of data availability) an MA(1) component can be present in the error term (see Evans, 1988, Hansen, Heaton and Yaron 1996). To assure consistent estimation, instruments must be used that are lagged at least twice.

We estimate (5) using annual data for Belgium (1958-1997). Note that, given that quarterly data are unavailable for Belgium, all estimations and simulations in this paper are based on relatively small samples containing annual data. These sample sizes are frequently used when estimating consumer Euler equations. Data and data sources are described in table 1. Note that for c_t we use real per capita consumer expenditures on non-durables and services [4]. For g_t we

[4] Using this series is based on the assumption that preferences over non-durables and the service flow of durables are additively separable. This is a commonly made assumption since Hall (1978).

use real per capita government consumption (wage and non-wage). Our conclusions are unaffected if instead we use aggregate government consumption (unreported). For r_t we use the real three month Treasury Bill rate. Kocherlakota (1990b) also uses this rate as a risk-free rate in annual data. It is important to note that this series reported by the IMF is a time average (see the discussion on time aggregation in section 3.1).

Table 1. Data and data sources.

c_t	Real per capita expenditures on non-durable goods and services. Calculated from several volumes from 'Statistisch Jaarboek van België' (NIS) for period 1958-1979. Calculated from 'Nationale Rekeningen' (INR, 1998) for period 1980-1997. Calculated as the difference between total conumption and expenditures on durables (including cars). Deflated by implicit price deflator for non-durables and services (1990=100).
g_t	Real per capita government consumption. Taken from OECD Economic Outlook (CD ROM 1999 Vol. 1) and reported in real terms with code CGV (deflated by implicit price deflator for government consumption, 1990=100).
r_t	Real three month Treasury Bill rate. Reported by IMF, International Financial Statistics, several volumes (line 60c). Available from 1958 onwards. Put into real terms by subtracting the inflation rate calculated from the implicit price deflator for non-durables and services.

Note: per capita measures are obtained after dividing by total population (from OECD Economic Outlook CD ROM 1999 Vol.1 with code POP).

We use the following instrument sets: instrument set 1 contains a constant, g(-2)/g(-3) and 1+r(-2). Instrument set 2 contains set 1 plus a dummy (DUM) which equals 1 in 1962 and 1963 since there is a significant outlier in g/g(-1) in these years. This is due to a political decision of the Belgian government to increase government real wages so as to lower the gap between public and private sector wages that had become substantial in the previous years. This decision was taken independently of the economic conditions (expected) in 1962-63 and was announced in the budget beforehand. This dummy thus belongs in the consumer's information set and is a valid instrument. Instrument set 3 contains a constant, c(-2)/c(-3), g(-2)/g(-3) and 1+r(-2). Instrument set 4 contains set 3 plus DUM. In instrument sets 1 and 2 c(-2)/c(-3) is left out since this variable is insignificant in the equations forecasting c/c(-1), g/g(-1) and r. The Akaike and Schwarz criteria of the forecasting equations for c/c(-1), g/g(-1) and r do not improve if an additional lag of each variable is added. For instrument set 1 there are no overidentifying restrictions to be tested since the number of parameters equals the number of instruments. For the other instrument sets the number of overidentifying restrictions is equal to the difference between the number of instruments and the number of estimated parameters.

The estimates and their t-statistics together with the test statistic for overidentifying restrictions and the adjusted R^2 for the OLS regression of c/c(-1), g/g(-1) and r on the instruments are presented in table 2. The test for overidentifying restrictions never rejects the model and the instrument set. Low values for this test may be due to the fact that some parameters are estimated imprecisely (in this case, θ - see below) which leads to a high variance of the sample moments (see Verbeek 1997).

Inspection of the errors does not exclude the possibility of first-order autocorrelation, so that the use of twice lagged instruments can be justified (not reported). The results with c(-2)/c(-3) as an instrument are not much different than the results without c(-2)/c(-3).

Table 2. Estimation of equation (5) using annual data for Belgium (1958-1997).

Instrument set	R^2 adj. of OLS regression on instruments			GMM estimation of eq. (5)			
	c/c(-1)	**g/g(-1)**	**r**	α	β	θ	**test(df)**
1	0.080	0.254	0.251	0.909 [0.051]	0.977 [0.415]	3.853 [2.537]	-
2	0.080	0.608	0.253	0.903 [0.039]	1.039 [0.081]	4.149 [1.916]	0.021 (1)
3	0.070	0.252	0.268	0.910 [0.046]	0.972 [0.409]	3.767 [2.203]	0.006 (1)
4	0.063	0.611	0.269	0.904 [0.033]	1.036 [0.063]	4.083 [1.544]	0.026 (2)

Notes: The statistics in columns 2,3 and 4 give the R^2adj of an OLS regression of c/c(-1), g/g(-1) and r on the instruments.The instruments contained in set 1 are a constant, g(-2)/g(-3) and 1+r(-2). Instrument set 2 contains set 1 plus a dummy which equals 1 in 1962 and 1963 (DUM). Instrument set 3 contains a constant, c(-2)/c(-3), g(-2)/g(-3) and r(-2). Set 4 contains set 3 plus DUM. The GMM estimates of the parameters in equation (5) are presented in columns 5,6 and 7 with Newey-West standard errors between square brackets (lag truncation = 3). Column 8 contains the chi-square test statistic for overidentifying restrictions, which is distributed chi-square with df = degrees of freedom equal to the number of instruments minus the number of estimated parameters.

The estimates for the discount factor (α), which reflects the subjective rate of time preference, are significant and have economically meaningful values, though they are rather low for a country like Belgium where the saving rate is relatively high. An explanation could be that this coefficient is biased. Ogaki and Reinhart (1998) argue that for US data there is a negative relationship between α and θ when estimating Euler equations. Thus if a 'high' estimate for θ is found (see below), it could be that the estimate for α is 'low'. So basically in this case, if one coefficient is poorly estimated, so will the other. We return to this issue in the next section.

The estimate of the coefficient of relative risk aversion $(1-\hat{\theta})$ or its inverse, the IES [5], is sometimes insignificant and has the wrong sign [6]. Thus the estimates for θ are implausible since they are outside concave parameter space. This is in line with Ogaki and Reinhart (1998), Holman (1998), Hansen and Singleton (1996) and Hall (1988). In the next section we assess whether the implausible estimates of θ can be attributed to the sample size used.

The point estimates for β are one or near to one for all instrument sets. The standard errors are much lower however when a dummy variable is used in the instrument set. Assuming these standard errors are well estimated, this could probably be attributed to the fact that the dummy variable is very significant in the forecasting equation for public consumption growth (i.e it is in that sense a good instrument). From these results, if we are interested in the possibility of an Edgeworth relationship, we would tend to conclude either that private and public consumption are independent or that the degree of dependence between them is low (since our

[5] If c and g are (Edgeworth) independent, then the IES of private consumption is equal to the IES of effective consumption.

[6] The asymptotic variance for the estimate of the IES can be calculated through $\sigma_{\theta}^2/(1-\hat{\theta})^4$.

measure $-U_{cg}/U_{cc}$ derived above will be close to zero if β is close to 1). The direction of the dependence is unknown, given the poor estimate of θ. The question is however whether the estimate for β and its estimated standard error are biased. If this is the case, this could suggest that the practice of taking a poor estimate of θ for granted and still draw inference on the other parameters in the model is in fact questionable. We investigate these issues further in the next section.

3. Monte Carlo simulations.

In this section we illustrate the possibility that the estimated coefficient of RRA / IES can be severely biased when estimated with GMM in a small sample. More specifically, we show that when we generate data under plausible positive values for the coefficient of RRA / IES, negative estimates can be obtained when using GMM.

Further, we look at whether the estimated parameter reflecting the potential relationship between private and government consumption is also biased. It is also important to investigate whether the standard errors on the parameters are well estimated.

3.1 Test setting.

Under the assumption that the growth rates of private consumption and government consumption and the logarithm of $1+r_{t+1}$ are jointly normally distributed, we can linearize equation (5) (see Hansen and Singleton, 1983). In appendix A2 we show that we can then write equation (5) as

$$(6)\ \hat{X}_{t+1} = \frac{\theta(1-\beta)}{1-\theta\beta}\hat{Z}_{t+1} + \frac{1}{1-\theta\beta}\hat{R}_{t+1} + \pi_{t+1}$$

where $\hat{X}_{t+1} \equiv \ln(c_{t+1}/c_t) - \bar{X}$, $\hat{Z}_{t+1} \equiv \ln(g_{t+1}/g_t) - \bar{Z}$ and $\hat{R}_{t+1} = \ln(1+r_{t+1}) - \bar{R}$. $\bar{X}$, $\bar{Z}$ and $\bar{R}$ are the sample means of $\ln(c_{t+1}/c_t)$, $\ln(g_{t+1}/g_t)$ and $\ln(1+r_{t+1})$ respectively. $\hat{X}_{t+1}$, $\hat{Z}_{t+1}$ and $\hat{R}_{t+1}$ are stationary variables [7]. The error term is given by $\pi_{t+1} = -1/(1-\theta\beta)v_{t+1}$ with $E_t v_{t+1} = 0$.

In section 2 it was argued that time aggregation might play a role in the estimation of the consumer Euler equation. Hansen, Heaton and Yaron (1996) show that if all variables in an equation like (6) are time-averages, the error term term v_{t+1} will have an MA(1) structure.

In one case - the limit case of continuous decision making (the number of decision periods within each observation period approaches infinity) - the first order autocorrelation coefficient of v_{t+1} will be $\rho = 0.25$ (see Hansen, Heaton and Yaron 1996 and Hall 1988). Since time aggregation imposes an MA(1) structure on the error v_{t+1} we can write $v_{t+1} = \xi_{t+1} + a\xi_t$ where ξ_t is i.i.d and normal with mean zero and constant variance. From $\rho = 0.25 = a/(1+a^2)$, we can derive a = 0.268 (see also Karras 1994).

Imposing different values for θ and β in (6) [8] and using the *observed* growth rate of real per capita government consumption (in deviation from its sample mean) for $\hat{Z}_{t+1}$ and the *observed* logarithm of $1+r_{t+1}$ (in deviation from its sample mean) for $\hat{R}_{t+1}$, while drawing ξ from a normal distribution [9], allows us to construct n artificial series $(\hat{X}_{t+1})_i$ (where i = 1, ... n) for each case considered. For all cases we set the number of replications equal to 5000 (n = 5000).

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[7] This is confirmed by the unit root tests in appendix B.

[8] No values for α are imposed (see appendix A2 where α was part of the constant of the linearized Euler equation).

[9] The variance of ξ is derived from the variance of the Euler equation error as obtained in section 2.

After the construction of $(\hat{X}_{t+1})_i$, these artificial series are again transformed by adding the sample mean X to obtain $(X_{t+1})_i$. We then construct $(c_{t+1}/c_t)_i = e^{(X_{t+1})_i}$.

For each generated sample i (containing 40 observations each) that consists of the observed series g_{t+1}/g_t and $1+r_{t+1}$ and the artificial series $(c_{t+1}/c_t)_i$ α, β and θ are estimated through the original non-linear Euler equation (5) with the iterative GMM procedure set out in section 2. For each case we then obtain 5000 estimates for α, θ and β.

We generate data assuming there is time aggregation. Under continuous decision making a = 0.268. If we assume there is no time aggregation we set a = 0. We use instrument sets 1 and 2 as defined in section 2. Using instrument sets 3 and 4 instead does not alter the conclusions (not reported).

3.2 Test results.

The results of the simulations are presented in table 3 (pages 14-16). In case 1 we generate data under the hypotheses β = 0.6 and θ = -3 (which coincides with a coefficient of relative risk aversion of 4) with instrument set 2 (including DUM) and a=0 (no time aggregation). Both β and θ are mean biased and median biased when estimated with GMM in a small sample. The bias in β is not very large while θ seems to be severely biased (the magnitude of the bias in both coefficients can be compared by normalising the bias with the standard deviation). Even though the data have been generated under θ = -3, the mean for θ is much larger (- 0.339).

Table 3. Monte Carlo results.

Notes: Bias = sample mean of the estimates minus the true parameter value. The 5% rejection rates rates for β and θ report the fraction of the 5 000 trials in which the true parameter lays outside 1.96 standard errors of the point estimate. The 5% level rejection rate test reports the fraction of the 5 000 trials in which the test statistic for overidentifying restrictions exceeds the 5% critical value of the chi-square distribution (this is 3.841 for cases with instrument set 2). T.A indicates cases where data is simulated with time aggregation (a = 0.268 - continuous decision making). Instrument sets 1 and 2 are defined in table 2.

Case 1: $\beta = 0.6$, $\theta = -3$, instrument set 2.			
	α	β	θ
Mean	1.007	0.611	-0.339
Median	1.009	0.595	-0.572
Standard deviation	0.105	0.524	4.322
Bias	-	0.011	2.661
5% level rejection rate β (%)		10.420	
5% level rejection rate θ (%)		50.120	
5% level rejection rate test (%)		0.004	

Case 2: $\beta = 0.6$, $\theta = -3$, T.A , instrument set 2.			
	α	β	θ
Mean	0.998	0.602	-0.032
Median	1.004	0.591	-0.402
Standard deviation	0.096	0.442	3.979
Bias	-	0.002	2.968
5% level rejection rate β (%)		10.960	
5% level rejection rate θ (%)		54.260	
5% level rejection rate test (%)		0.004	

Case 3: $\beta = 0.8$, $\theta = -3$, instrument set 2.			
	α	β	θ
Mean	1.008	0.794	-0.395
Median	1.009	0.736	-0.636
Standard deviation	0.106	0.912	4.440
Bias	-	-0.006	2.605
5% level rejection rate β (%)		29.240	
5% level rejection rate θ (%)		48.400	
5% level rejection rate test (%)		0.004	

Case 4: $\beta = 0.8$, $\theta = -3$, T.A , instrument set 2.			
	α	β	θ
Mean	0.999	0.772	-0.048
Median	1.003	0.718	-0.383
Standard deviation	0.108	0.964	4.472
Bias	-	-0.028	2.952
5% level rejection rate β (%)		29.540	
5% level rejection rate θ (%)		53.180	
5% level rejection rate test (%)		0.002	

Case 5: $\beta = 1$, $\theta = -3$, instrument set 2.			
	α	β	θ
Mean	1.006	0.995	-0.364
Median	1.010	0.929	-0.671
Standard deviation	0.107	1.575	4.598
Bias	-	-0.005	2.636
5% level rejection rate β (%)		26.760	
5% level rejection rate θ (%)		48.080	
5% level rejection rate test (%)		0.002	

Table 3 (continued).

Case 6: $\beta = 1$, $\theta = -3$, T.A , instrument set 2.			
	α	β	θ
Mean	1.001	0.903	-0.156
Median	1.005	0.890	-0.505
Standard deviation	0.103	1.269	4.294
Bias	-	-0.097	2.844
5% level rejection rate β (%)		29.920	
5% level rejection rate θ (%)		52.600	
5% level rejection rate test (%)		0.002	

Case 7: $\beta = 0.6$, $\theta = -5$, instrument set 2.			
	α	β	θ
Mean	1.031	0.624	-1.147
Median	1.034	0.599	-1.549
Standard deviation	0.139	0.303	5.631
Bias	-	0.024	3.853
5% level rejection rate β (%)		15.140	
5% level rejection rate θ (%)		51.440	
5% level rejection rate test (%)		0.000	

Case 8: $\beta = 0.6$, $\theta = -5$, T.A , instrument set 2.			
	α	β	θ
Mean	1.020	0.609	-0.741
Median	1.026	0.595	-1.247
Standard deviation	0.134	0.278	5.447
Bias	-	0.009	4.259
5% level rejection rate β (%)		16.600	
5% level rejection rate θ (%)		55.380	
5% level rejection rate test (%)		0.000	

Case 9: $\beta = 0.6$, $\theta = -1$, instrument set 2.			
	α	β	θ
Mean	0.979	0.576	0.648
Median	0.981	0.578	0.553
Standard deviation	0.059	0.572	2.517
Bias	-	-0.024	1.648
5% level rejection rate β (%)		5.700	
5% level rejection rate θ (%)		57.720	
5% level rejection rate test (%)		0.008	

Case 10: $\beta = 0.6$, $\theta = -1$, T.A , instrument set 2.			
	α	β	θ
Mean	0.972	0.598	0.922
Median	0.977	0.592	0.706
Standard deviation	0.058	0.590	2.483
Bias	-	-0.002	1.922
5% level rejection rate β (%)		7.760	
5% level rejection rate θ (%)		62.900	
5% level rejection rate test (%)		0.001	

Table 3 (continued).

Case 11: $\beta = 0.6$, $\theta = 0.5$, instrument set 2.			
	α	β	θ
Mean	0.963	0.632	1.270
Median	0.966	0.599	1.175
Standard deviation	0.027	0.487	1.076
Bias	-	0.032	0.770
5% level rejection rate β (%)		9.420	
5% level rejection rate θ (%)		55.500	
5% level rejection rate test (%)		0.000	

Case 12: $\beta = 0.6$, $\theta = 0.5$, T.A , instrument set 2.			
	α	β	θ
Mean	0.962	0.651	1.333
Median	0.964	0.609	1.220
Standard deviation	0.030	0.509	1.108
Bias	-	0.051	0.833
5% level rejection rate β (%)		12.02	
5% level rejection rate θ (%)		57.84	
5% level rejection rate test (%)		0.000	

Case 13: $\beta = 0.6$, $\theta = -3$, instrument set 1.			
	α	β	θ
Mean	0.976	0.618	1.245
Median	0.987	0.516	0.370
Standard deviation	0.163	4.039	6.716
Bias	-	0.018	4.245
5% level rejection rate β (%)		12.26	
5% level rejection rate θ (%)		28.38	
5% level rejection rate test (%)		-	

Case 14: $\beta = 0.6$, $\theta = -3$, T.A , instrument set 1.			
	α	β	θ
Mean	0.971	0.499	1.427
Median	0.983	0.509	0.512
Standard deviation	0.155	4.290	6.363
Bias	-	-0.101	4.427
5% level rejection rate β (%)		12.54	
5% level rejection rate θ (%)		32.02	
5% level rejection rate test (%)		-	

Case 15: $\beta = 1$, $\theta = -3$, instrument set 1.			
	α	β	θ
Mean	0.975	0.856	1.238
Median	0.983	0.711	0.441
Standard deviation	0.155	4.808	6.745
Bias	-	-0.144	4.238
5% level rejection rate β (%)		21.36	
5% level rejection rate θ (%)		30.66	
5% level rejection rate test (%)		-	

Case 16: $\beta = 1$, $\theta = -3$, T.A , instrument set 1.			
	α	β	θ
Mean	0.979	0.875	1.035
Median	0.986	0.728	0.333
Standard deviation	0.154	7.863	6.605
Bias	-	-0.125	4.035
5% level rejection rate β (%)		20.64	
5% level rejection rate θ (%)		29.04	
5% level rejection rate test (%)		-	

Thus, estimates of θ are indeed poor with GMM in a small sample; they are upward biased. As can be seen from the standard deviaton, there is a considerable dispersion in the distribution of θ.

Standard inference is based on the assumption that the t-values are asymptotically standard normally distributed, while the test for overidentifying restrictions is asymptotically chi-square distributed. We look at whether using this assumption in a small sample gives rise to poor inference, i.e do the small sample distributions of these test statistics approximate their asymptotic counterparts. Table 3 reports the fraction of the 5000 trials in which the null hypotheses for β and θ were rejected using the 5% critical values of the standard normal distribution. The rejection rate for the test statistic for overidentifying restrictions is also reported. Since the null hypotheses are true (by construction), the rejection rates (i.e type I errors) should be close to the significance level of 5%.

In case 1 (table 3) we can see that the rejection rates are quite high. For β we reject the null hypothesis in 10 % of the cases despite the fact that we know that it is true. For θ this percentage is even 50 %. This is probably due to the severity of the bias in the parameters and the underestimation of their asymptotic standard errors. Inference on β is more reliable than inference on θ.

In case 2 (table 3) time aggregation is assumed to be present. The results are very similar to those in case 1. If we compare the other cases with and without time aggregation the results are not very different, except perhaps that the bias in the estimate of θ tends to be somewhat larger if time aggregation is present (except in case 15-16).

Keeping the value of θ fixed, we can see what happens for different values of β by looking at cases 3,4,5 and 6. For cases without time aggregation, the bias in the estimate of β remains very small. In cases 4 and 6 however there appears to be a larger downward bias in the estimate of β. From the standard deviations it also seems that β is measured less precisely if its true value approaches 1, both in cases with and without time aggregation. Moreover the 5% rejection rate also rises as β increases, suggesting that inference is problematic in these cases. The estimates, the standard deviations and the 5% rejection rates for θ do not vary much if the true value of β is changed.

In cases 7, 8, 9, 10, 11 and 12 we change the true value of θ while keeping the value of β fixed at 0.6 . In cases 7 and 8 for instance we set $\theta = -5$ [10]. This lowers the estimates of θ compared to cases 1 and 2. This does not seem to affect the estimates of β too much. This is also the case when θ is set to -1 or 0.5. Thus the estimates of β seem to be independent of the true value of θ. This conclusion holds for other values of β as well (not reported).

In cases 13,14,15 and 16 we work with instrument set 1 which does not include DUM. The result is that the bias in the estimates of both β and θ is larger (compare cases 13,14 with 1,2 and 15,16 with 3,4). The standard deviation of the β distribution becomes much larger. Again the normal distribution does not seem to be very good to test hypotheses given the large rejection rates, especially when the true value of β approaches 1.

In general we can say that the mean bias in the estimate of β never becomes very large and is always much smaller than for θ. The rejection rates at the 5% level for β are between 5.7 and

[10] Kocherlakota (1990b) even justifies setting $\theta = -12$ (a coefficient of RRA as large as 13).

29.9%. Given the relatively small bias in the estimate for β this overrejection could indicate that in some cases standard errors are seriously understated. In other cases the precise estimation of and correct inference on β using the normal distribution seems to be possible in a small sample. The normal distribution is never a good approximation for the small sample distribution of θ.

For our estimations of β in section 2 these results could imply that β is not biased very much (so that the conclusion of low dependence between c_t and g_t remains valid as far as the point estimates are concerned), but that the asymptotic standard errors on β might be underestimated.

The test statistic for overidentifying restrictions tends to under-reject the model in all 12 first cases where it can be calculated. This finding could explain the low values found for this test in section 2. It seems that this test is not very reliable in small samples to test the overidentifying restrictions. The finding of under-rejection is not in accordance with Mao (1990), who finds over-rejection. Tauchen (1986) and Ferson and Foerster (1994), on the other hand, tend to find under-rejection. Besides sample size, additional reasons for the disparity between the small sample distribution and the asymptotic distribution of the test for overidentifying restrictions have been put forward. Tauchen (1986) finds that using too many lags to form instruments leads to under-rejection in small samples. Ferson and Foerster (1994) find that with small sample sizes the use of *iterated* GMM leads to an under-rejection of their model while using *two-step* GMM they tend to over-reject their model. Mao (1990) does not find significant differences between both methods as far as the test for overidentifying restrictions is concerned.

The test setting where data is generated through (6) is convenient because, by construction, one parameter (α) is left out of the analysis so that no hypotheses for α must be imposed in the simulations. As noted in section 2, there may be a negative relationship between α and θ when estimating consumer Euler equations. This seems to be confirmed by the simulation results in table 3 where the larger the estimates of θ, the more the estimate for α is below one. This could explain the low values found for α in section 2. Values for α larger than one are frequently found in empirical studies using quarterly data (van Dalen 1999, Ni 1995, ...) where the coefficient of relative risk aversion is well estimated. These values can be justified on a theoretical basis (for the intuition of $\alpha > 1$ in growing economies, see Kocherlakota 1990a). As argued by van Dalen (1999), however, they might as well be an artefact of macroeconomic data (i.e due to aggregation issues).

The results of the simulations can help to explain some of the findings in section 2, they are not usable for inference however. The reason is that the data for the simulations are constructed under certain assumptions which may not hold in practice. The moment conditions may well be generated by non-normal data generating mechanisms. Moreover, the case of time-aggregation considered is a limit case. An alternative approach is to create a bootstrap simulation based on *actual* data. In appendix C we report the results of a bootstrap simulation conducted through the resampling of residuals obtained from the estimation of (5). To capture the potential autocorrelation in the error (due to time aggregation) we have also resampled blocks of residuals.

This exercise makes it possible to check the *robustness* of some of the results obtained with the Monte Carlo simulations. It also makes *inference* possible through the calculation of bootstrapped standard errors and through the simulation of small sample distributions to test hypotheses on the parameters of interest. The main conclusions of these bootstrap simulations are that - when generating data under plausible (negative) values for θ - we find that the estimates of θ tend to have a severe upward bias. The estimates for β are in general not as seriously biased. Further, using the rejection rates (defined as before), we find that the normal distribution is not a good guide for inference on the parameters. Finally, there is under-rejection of the model using the test statistic for overidentifying restrictions. These conclusions thus tend to confirm the earlier findings when data was generated under the normality assumption.

As far as inference is concerned, we also have calculated bootstrapped standard errors for all three parameters using the estimated values reported in section 2. For all parameters the bootstrapped standard errors are sometimes in the same range but often higher than the asymptotic (Newey-West) standard errors reported in table 2. This confirms what we suspected: the asymptotic standard errors may be underestimated.
The values of the bootstrapped standard errors for β are such that, using the critical values of the standard normal distribution, we are not able to stick to the conclusions found in section 2 (we cannot reject β=0 for instance). However, as noted before, the normal distribution is a poor approximation to the small sample distribution. For hypothesis testing on the parameters of interest the use of bootstrapped distributions may then be a better alternative. Using the simulated small sample distributions for β, for instance, does allow us to reject $\beta = 0$ but not $\beta = 1$.

4. Conclusions.

Many researchers question the feasibility of estimating the coefficient of relative risk aversion using macroeconomic data. In practice, the estimation of the parameters of consumer Euler equations (among which is the coefficient of relative risk aversion) is still widely applied.

Some studies find plausible estimates for the coefficient of relative risk aversion, others do not. In this paper we investigate whether the often found negative estimates for this coefficient could at least partially be attributed to the sample size used. We have indications that this could be the case. Moreover, we want to know whether other parameters in the Euler equation can be well estimated and whether inference on other parameters is possible given a poorly estimated coefficient of relative risk aversion. From Monte Carlo simulations applied to an example where a consumer Euler equation is estimated using annual data for Belgium, we find that not all parameters are necessarily as severely biased. For some parameters inference, using the normal distribution as an approximation to the small sample distribution, is more reliable than for others. However this is only true in certain cases, depending among other things on the true values of these parameters. In general, unless one is willing to perform a bootstrap simulation, using a sample size that is too small should be avoided and finding a poorly estimated coefficient of relative risk aversion must be interpreted as a signal that there might be something wrong with the other estimated parameters as well.

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Appendix A1 : derivation of (3).

The maximization of (1) occurs subject to the budget constraint, (A1.1) $(1+r_j)(y_{j-1}-c_{j-1}+w_{j-1})=w_j$

where r_j is the interest rate, w_j is financial wealth at the beginning of period j, y_{j-1} is disposable labour income and c_{j-1} are consumer expenditures. We can write the value function for j = t as,

(A1.2) $V_t(w_t)=\max_{w_{t+1}}\left[U(c_t,g_t)+\alpha E_t V_{t+1}(w_{t+1})\right]$

where w_t is the state variable defined by the budget constraint (A1.1). We maximize the term between brackets on the RHS of (A1.2) with respect to c_t and subject to (A1.1) written for j=t+1. This leads to

(A1.3) $\dfrac{\partial U}{\partial c_t}-\alpha E_t\{V'(w_{t+1})(1+r_{t+1})\}=0$

By the envelope theorem we know that $V'(w_{t+1})=\dfrac{\partial U}{\partial c_{t+1}}$, so we obtain equation (2) in the text.

Appendix A2 : derivation of (6).

Suppose $x_{t+1}\equiv c_{t+1}/c_t$ and $z_{t+1}\equiv g_{t+1}/g_t$ and write $q_{t+1}\equiv x_{t+1}^{\theta\beta-1}z_{t+1}^{\theta(1-\beta)}(1+r_{t+1})$. Then (5) can be written as

(A2.1) $E_t(q_{t+1})=1/\alpha$

Write $X_{t+1} \equiv \ln x_{t+1}$, $Z_{t+1} \equiv \ln z_{t+1}$, $R_{t+1} \equiv \ln(1+r_{t+1})$ and $U_{t+1} \equiv \ln q_{t+1}$. Assume now that X, Z and R are stationary, jointly normally distributed variables. Then the distribution of U_{t+1} conditional on information at time t is normal with constant variance σ^2 and mean μ. By the lognormal property [11] we can then write,

$$(A2.2)\quad E_t(q_{t+1}) = \exp\left(\mu + \sigma^2/2\right)$$

From (A2.1) and (A2.2) we obtain,

$$(A2.3)\quad \exp(\mu + \sigma^2/2) = 1/\alpha$$

We can write this as

$$(A2.4)\quad \mu = -\ln\alpha - \sigma^2/2$$

Now define a variable $v_{t+1} \equiv U_{t+1} - \mu$. From the definitions of U_{t+1} and q_{t+1} and from (A2.4) we can write,

$$(A2.5)\quad v_{t+1} = (\theta\beta - 1)X_{t+1} + \theta(1-\beta)Z_{t+1} + R_{t+1} + \ln\alpha + \sigma^2/2$$

[11] The lognormal property says that if y is a normal variable with mean E(y) and variance σ^2 we can write, $E(\exp(y)) = \exp(E(y) + \sigma^2/2)$.

with $E_t(v_{t+1}) = 0$. If we write the variables X_{t+1}, Z_{t+1} and R_{t+1} as deviations from their sample mean $\hat{X}_{t+1}$, $\hat{Z}_{t+1}$ and $\hat{R}_{t+1}$, we can drop the constant $\log\alpha + \sigma^2/2$. (A2.5) becomes,

$$\text{(A2.6)} \quad v_{t+1} = (\theta\beta - 1)\hat{X}_{t+1} + \theta(1-\beta)\hat{Z}_{t+1} + \hat{R}_{t+1}$$

Solving for $\hat{X}_{t+1}$ results in equation (6).

Appendix B: unit root tests.

Table B1. Augmented Dickey-Fuller unit root tests.

Sample 1958-1997			
Series	Lags	Trend	No Trend
c	4	-2.313	-0.002
g	2	-0.698	-2.527
r	0	-3.585**	-3.248**
c/c(-1)	0	-4.402***	-4.136***
g/g(-1)	0	-6.369***	-3.356**

Note: all regressions include a constant. The entries in the columns 'trend' and 'no trend' denote the DF test statistic if a deterministic trend is or is not included in the regression; the test statistics are compared to the MacKinnon critical values for the rejection of the null hypothesis of a unit root. *** denotes significance at the 1% level, ** denotes significance at the 5% level, * denotes significance at the 10% level. 'Lags' denotes the number of lags included to remove serial correlation from the residuals. Starting with 4 lags, the number of lags has been reduced based on standard t-tests and the Akaike information criterion.

c = real per capita private consumption of non-durables and services.

g = real per capita government consumption (wage and nonwage).

r = the 3-month Treasury Bill rate in real terms.

In section 2 we work with the variables r, c/c(-1) and g/g(-1). In section 3 we work with the logarithms of c/c(-1), g/g(-1)) and 1+r : the conclusions concerning unit roots are exactly the same as with c/c(-1), g/g(-1) and r.

Appendix C: bootstrap simulation - method and results.

1. Method.

To implement the bootstrap we proceed as follows,

1) We estimate equation (5) - restricted or unrestricted - which gives a series of residuals. We consider the following cases:

- case 1: different restrictions are imposed on β and θ (as in the Monte Carlo simulations). We report the results for a) $\beta = 0.6$, $\theta = -3$ (instrument set 2), b) $\beta = 0.6$, $\theta = -3$ (instrument set 1) and c) $\beta = 1$, $\theta = -3$ (instrument set 2).
- case 2: no restrictions are imposed on the parameters (as in section 2).
- case 3: restrictions are imposed on β, a) $\beta = 0$ and b) $\beta = 1$.

2) We randomly draw 40 of the obtained residuals (with replacement). This gives a pseudo time-series of residuals.

3) Using the actual g_{t+1}/g_t and r_{t+1} series together with the resampled residuals and setting the values of α, β and θ to the values imposed or estimated in cases 1,2 and 3, we construct an artificial series c_{t+1}/c_t (containing 40 observations) using the fact that the Euler equation error ε_{t+1} is given by $\varepsilon_{t+1} = \alpha(1+r_{t+1})(c_{t+1}/c_t)^{\theta\beta-1}(g_{t+1}/g_t)^{\theta(1-\beta)} - 1$.

4) We repeat steps 2 and 3 n times. Thus we construct n artificial series $(c_{t+1}/c_t)_i$ (i=1,...,n) for each case considered. For all cases we set n = 5000. For each generated

sample i (containing 40 observations each) that consists of the observed series g_{t+1}/g_t and $1+r_{t+1}$ and the artificial series $(c_{t+1}/c_t)_i$ α , β and θ are estimated through the original non-linear Euler equation (5) with the iterative GMM procedure set out in section 2. For each case we then obtain 5000 estimates for α, θ and β from which we calculate bootstrap distributions.

Case 1 (a,b,c) is similar to some cases in table 3 of the paper and can serve as a test to check the robustness of the Monte Carlo simulations. We can check whether we also find biases in the point estimates and whether inference using the standard normal distribution is reliable. Unlike with the Monte Carlo simulations, we do not compare different cases.

Starting from the estimated values of the parameters (as reported in table 2 of the paper) case 2 allows to calculate bootstrapped standard errors which can be compared to the asymptotic (Newey-West) standard errors reported in table 2. The bootstrapped standard error of a parameter estimate is calculated as the standard deviation of the 5000 bootstrapped estimates of that parameter.

Case 3 allows us to simulate the small sample distribution of β and of its t-values under the hypotheses $\beta = 0$ and $\beta = 1$. The use of the bootstrapped distribution of t-ratios replaces the standard normal distribution to test the hypotheses $\beta = 0$ and $\beta = 1$.

To take into account the fact that the error term may be autocorrelated due to time aggregation we have also resampled blocks of residuals. This makes sense given that the error term structure is known but the MA(1) parameter is not (unless in the limit case of continuous decision making). We use the resampling method as discussed by Li and Maddala (1996) where overlapping blocks of residuals are estimated with replacement [12]. Given the fact that the conclusions are identical to those of the simulations where individual observations are resampled, we do not report the results of these simulations.

2. Results.

In table C1 we report the results of the bootstrap simulations of case 1 (a,b,c). The results point in the same direction as the findings of the Monte Carlo simulations. When generating data under plausible (negative) values for θ, we find that the estimates of θ tend to have a severe upward bias. The estimates for β are in general not as seriously biased. The bias in β seems to depend on its true value[13]. Further, using the rejection rates (defined as before), we find that the normal distribution is not a good guide for inference on the parameters (except perhaps in case 1a for β). Finally, there is under-rejection of the model using the test statistic for overidentifying restrictions.

Table C1. Bootstrap results: case 1 (a,b,c).

Case 1a : $\beta = 0.6$, $\theta = -3$, instrument set 2			
	α	β	θ
Mean	0.981	0.599	0.968
Median	0.991	0.601	0.538
Standard deviation	0.103	0.623	3.227
Bias	- 0.139	-0.001	3.968
5% level rejection rate β (%)		7.300	
5% level rejection rate θ (%)		68.380	
5% level rejection rate test		0.000	

[12] See Li and Maddala (1996) for the advantages of this method. The choice of the length of the blocks depends on the form of the error process, the sample size but can also be based on subjective judgement (sample correlations). Given the theoretical possibility of an MA(1) error, Li and Maddala (1996) suggest a block length of 2 or 3 for a sample of size 200. Given the small number of observations in our sample and given that block length should increase with sample size, we choose block length = 2.

[13] Note that the value imposed for α to generate data is its estimated value (under the restrictions imposed on β and θ). From table C1 (bias) we can see that the mean of the bootstrapped α's is lower than the 'true' value for α. This seems to offer support for the conclusion in section 3 of the paper that 'large' estimates for θ tend to coincide with 'small' estimates for α.

Table C1 (continued).

Case 1b : $\beta = 0.6$, $\theta = -3$, instrument set 1			
	α	β	θ
Mean	0.943	0.478	2.367
Median	0.973	0.478	1.278
Standard deviation	0.126	3.215	4.247
Bias	- 0.175	-0.122	5.367
5% level rejection rate β (%)		14.000	
5% level rejection rate θ (%)		46.840	
5% level rejection rate test		-	

Case 1c : $\beta = 1$, $\theta = -3$, instrument set 2			
	α	β	θ
Mean	0.996	0.764	0.451
Median	0.999	0.799	0.204
Standard deviation	0.115	1.306	3.697
Bias	- 0.109	-0.236	3.451
5% level rejection rate β (%)		33.420	
5% level rejection rate θ (%)		58.340	
5% level rejection rate test		0.000	

Notes: Bias = sample mean of the estimates minus the true parameter value. The 5% rejection rates rates for β and θ report the fraction of the 5 000 trials in which the true parameter lays outside 1.96 standard errors of the point estimate. The 5% level rejection rate test reports the fraction of the 5 000 trials in which the test statistic for overidentifying restrictions exceeds the 5% critical value of the chi-square distribution (this is 3.841 for cases with instrument set 2). Instrument sets 1 and 2 are defined in table 2.

As far as inference is concerned, we also have calculated bootstrapped standard errors for all three parameters using the estimated values reported in section 2 of the paper. These are reported in table C2. The bootstrapped standard errors are sometimes in the same range but often higher than the asymptotic (Newey-West) standard errors reported in table 2 of the paper. This confirms what we suspected: the asymptotic standard errors may be underestimated.

The values of the bootstrapped standard errors are such that, using the critical values of the standard normal distribution for β, we are not able to stick to the conclusions found in section 2 (we cannot reject β=0 for instance). However, as noted before, the normal distribution may be a poor approximation to the small sample distribution.

Table C2. Bootstrapped standard errors.

	α	β	θ
Instrument set 1	0.050	3.276	2.330
Instrument set 2	0.043	0.805	1.989

For hypothesis testing on the parameters of interest the use of bootstrapped distributions may be a better alternative. Using the critical values of the simulated small sample distributions of the t-ratios of β [14], for instance, does allow us to reject $\beta = 0$ but not $\beta = 1$.

[14] Available upon request.

Imperfect information and the excess sensitivity of private consumption to government expenditures.

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Abstract.

In this paper we consider a new explanation for the often encountered observation that private consumption is excessively sensitive to anticipated government expenditures. We show that this excess sensitivity arises if consumers are aware of the government's intertemporal budget constraint, but lack information on the aggregate economy. Given the strong assumption that consumers incorporate the government budget constraint, we test our model in three high debt countries where it is more likely that consumers have developed an awareness for government issues. In some of these countries and especially during periods of high debt accumulation, we observe some excess sensitivity with respect to (lagged) income *and* government expenditures which can be interpreted as evidence supporting our model.

JEL Classification: E62, E21, D91.

Keywords: private consumption, government expenditures, excess sensitivity, government budget constraint, imperfect information.

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1.Introduction.

The permanent income hypothesis implies that changes in aggregate private consumption are unpredictable (see Hall 1978). In reality, however, private consumption changes are generally found to be 'excessively sensitive' to predictable changes in *income* and are sometimes found to be 'excessively sensitive' to predictable changes in *government expenditures*.

Excess sensitivity of private consumption with respect to *income* has been interpreted most often by assuming the presence of liquidity constrained consumers (see Campbell and Mankiw 1990, 1991). In each period these consumers consume their entire disposable income. Since they cannot save nor borrow, this implies that they can only change their consumption when changes in income effectively materialize. Thus the possibility emerges that consumption changes do occur in response to previously anticipated changes in income.

Other authors have come up with alternative explanations for the failure of Hall's random walk hypothesis of private consumption at the aggregate level. Gali (1990) considers finite lifetimes, whereas Goodfriend (1992), Pischke (1995) and Demery and Duck (2000) assume consumers who have imperfect information on aggregate variables.

Excess sensitivity of private consumption to *government expenditures* has been interpreted in terms of an Edgeworth relationship between private consumption and government expenditures. Aschauer (1985) has modified Hall's Euler equation to allow for (Edgeworth) substitutability or complementarity effects between private consumption and government expenditures. Existing studies have found different results for different countries (Karras 1994, Evans and Karras 1996, Ni 1995). To some extent this is not surprising. One may indeed be sceptical about the validity of an Edgeworth relationship at the aggregate level since some components of government expenditures are likely to be complements to private consumption, while others are likely to be substitutes. Evans and Karras (1998), for instance, show for a sample of 66 countries that private consumption and non-military government spending are generally substitutes or independent, whereas private consumption and military spending are better described as complements.

In this paper we focus on the idea of imperfect information as developed by Pischke (1995) and Demery and Duck (2000). Imperfect information can explain why, in contradiction to the permanent income hypothesis, private consumption is observed to be both 'excessively smooth' and 'excessively sensitive' to income (the so-called 'Deaton paradox'). Moreover, as an innovation of this paper, assuming imperfectly informed consumers who incorporate the government budget constraint provides an alternative interpretation for the observed excess sensitivity of private consumption to government expenditures at the aggregate level.

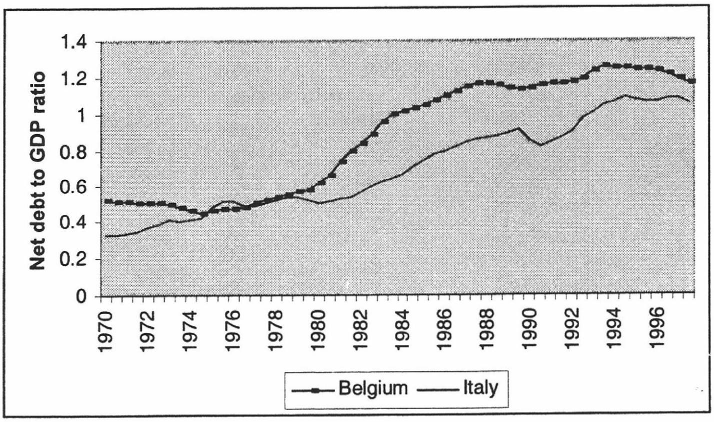
The failure of the random walk hypothesis of private consumption at the aggregate level can occur if individual income and/or the individual's perception of government expenditures has an aggregate and an individual-specific component. Consumers cannot differentiate between both components because they do not know the aggregate component. If there is an innovation in aggregate income or in government expenditures, this innovation will be partly misinterpreted as an innovation in the individual-specific component. Given that the aggregate component is more persistent than the individual-specific component, permanent income will not be adjusted appropriately and consumption will be too smooth. In the next period(s), consumers will notice that the change in income or government expenditures persists and will adjust consumption again so that it will appear excessively sensitive to changes in income and/or government expenditures.

We consider a model with utility maximizing permanent income consumers who 1) are imperfectly informed because they do not observe aggregate variables and 2) are aware that the government must respect a budget constraint. As far as the first assumption is concerned, Pischke (1995) argues that consumers in the US may indeed be imperfectly informed about aggregate variables and may have little incentive to obtain this information since calculations suggest that the benefits of obtaining aggregate information are rather small compared to the costs. As for the second assumption, Lopez et. al. (2000) mention that assuming that consumers internalize the government budget constraint 'imposes formidable requirements on agents' ability to gather and process information'. Especially in our case, where agents are not perfectly informed, this assumption requires a justification. We believe that the level and increase of the government debt in the economy may play a crucial role. It is likely that consumers in high-debt countries will benefit relatively more from information on government financing issues than consumers in countries without a problematic debt history. They are therefore more likely to develop an awareness for the government budget constraint (even if they have no incentive to obtain or to use exact aggregate information). Some authors have noted that consumers in high-debt countries tend to be more aware of the government budget constraint (see for instance Nicoletti 1988, 1992; Dalamagas 1993). The knowledge of the future tax implications of debt may have a higher value for consumers in high-debt countries because they may feel that the 'day of reckoning' is imminent. Assuming that the idea of imperfectly informed consumers incorporating the government budget constraint is relevant to explain excess sensitivity of private consumption with respect to government expenditures, we would especially expect it to hold in countries characterized by a problematic debt situation[1]. Therefore we estimate our model for a number of OECD countries (Belgium, Italy and Greece) where the debt to GDP ratio has increased rapidly since

[1] Though there is no reason why it could not hold in low-debt countries as well.

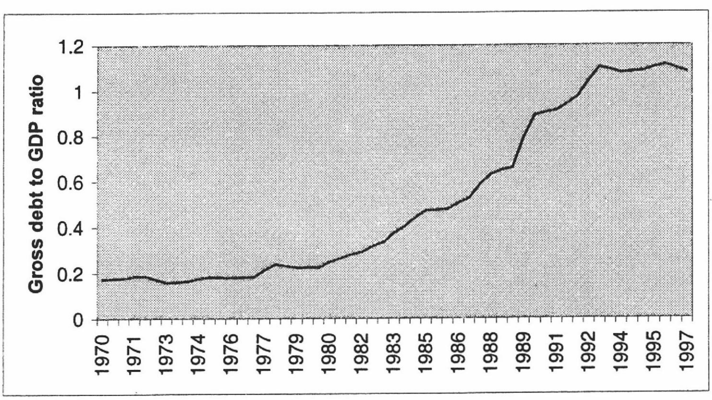
the mid-seventies. This can be seen in figure 1 for Italy and Belgium where the *net* debt to GDP ratio has reached very high levels, even exceeding 1 during the nineties. In figure 2 we present the evolution of the *gross* debt to GDP ratio for Greece (for which no data on the net debt are available). We can see that the Greek debt situation has become problematic especially since the early eighties (exceeding 50% during the mid-eighties).

Figure 1. Net debt to GDP ratio in Belgium and Italy (1970:01 – 1997:02).



Source: OECD (Economic Outlook CD ROM 2001 Vol. 2).

Figure 2. Gross debt to GDP ratio in Greece (1970:01 – 1997:02).



Source: OECD (Economic Outlook CD ROM 2001 Vol.2).

The remainder of this paper is as follows. In section 2 we present our model which generates the possibility that private consumption changes are excessively sensitive to (lagged) pre-tax income and government expenditure changes. In section 3 we discuss an empirically testable consumption function as well as the data that we use and a number of methodological issues. In section 3 we also present our empirical results. Section 4 concludes.

2. The Model.

We consider an economy with the following characteristics:

1) The economy is populated by a very large number (n) of infinitely lived utility maximizing permanent income consumers. Each has a quadratic utility function[2]. We assume that the subjective rate of time preference for all consumers is equal to the constant real interest rate in the economy.

2) The macro structure of the economy is given by the following three equations which describe the processes for aggregate per capita pre-tax income and government expenditures and the intertemporal government budget constraint,

(1) $y_t = y_t^p + \varepsilon_t^y$

(2) $g_t = g_t^p + \varepsilon_t^g$

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[2] Assuming a utility function of the constant relative risk aversion type will introduce precaution. Though no closed-form solution exists in this case without further assumptions, precaution is then usually captured by allowing the consumer's discount rate for future disposable income to be larger than the discount rate of the government (i.e the real interest rate) with the difference reflecting a risk premium (see Muellbauer and Lattimore 1995). This is equivalent to the assumption of finite horizons (see Blanchard 1985). In these cases the incorporation of the government intertemporal budget constraint will be incomplete. In this paper we choose to focus on consumers who do incorporate the government budget fully but who have incomplete information sets.

$$(3)\ \sum_{j=0}^{\infty}(1+r)^{-j}g_{t+j}+b_{t-1}=\sum_{j=0}^{\infty}(1+r)^{-j}t_{t+j}$$

In these equations y_t is per capita pre-tax income, g_t is per capita government expenditures, t_t is per capita net taxes, r is the exogenous[3] real interest rate and b_{t-1} is per capita government debt. The latter is measured at the end of period t-1. Eq. (3) is the intertemporal budget constraint of the government. We assume that the government does not engage in Ponzi games so that $\lim_{j\to\infty}(1+r)^{-j}b_{t+j-1}=0$ holds. The intertemporal budget is balanced. In (1) and (2) we assume that pre-tax income and government expenditures can be written as the sum of a temporary component (white noise terms ε_t^y and ε_t^g respectively) and a permanent component. The latter components are random walks given by,

$$(4)\ \ y_t^P = y + y_{t-1}^P + v_t^y$$

$$(5)\ \ g_t^P = g + g_{t-1}^P + v_t^g$$

where v_t^y and v_t^g are white noise terms (see for instance Deaton 1992). Note that our results are not dependent upon our assumption of the processes for pre-tax income and government expenditures (see below).

3) At the micro level, pre-tax income and net taxes of each consumer i, y_{it} and t_{it} can be written as the sum of aggregate per capita pre-tax income and net taxes respectively and an individual-specific component. Furthermore, imperfectly informed consumers are assumed to

3 We assume that the real interest rate is determined on the international capital market and is thus not influenced by (peoples' expectations about) the level of government borrowing.

observe per capita government expenditures with noise (see below). This is captured by the following equations,

(6) $y_{it} = y_t + \varepsilon_{it}^y$

(7) $t_{it} = t_t + \varepsilon_{it}^t$

(8) $g_{it} = g_t + \varepsilon_{it}^g$

where ε_{it}^y , ε_{it}^t and ε_{it}^g are individual-specific white noise terms. These terms are independently distributed across all consumers. Their variance is constant across consumers.

4) If consumers were perfectly informed they would observe their own past consumption and their own current and past income and taxes. They would observe government expenditures without noise, thus $\varepsilon_{it}^g = 0$. They would also observe past aggregate per capita consumption and current and past aggregate per capita income, taxes and government expenditures (both permanent and transitory components). At the end of period t, when deciding on consumption, perfectly informed consumers would have information set $I_{it}^{PE} = \{c_{it-1}, c_{it-2}, ..., x_{it}, x_{it-1}, ..., c_{t-1}, c_{t-2}, ..., x_t, x_{t-1}, ...\}$ where c_{it-1} is i's consumption in period t-1 and c_{t-1} is total per capita consumption in period t-1. x_t is a three-dimensional vector (y_t, t_t, g_t) and x_{it} is the vector (y_{it}, t_{it}, g_{it}) [4].

In this paper we assume that all consumers are imperfectly informed about the aggregate economy. At the end of period t, when deciding on consumption, imperfectly informed consumers have information set $I_{it}^{IM} = I_{it}^{PE} \setminus \{c_{t-1}, c_{t-2} ..., x_t, x_{t-1}, ...\}$. We make the additional

[4] Note that b_t, b_{t-1},... is also a part of the information set of a perfectly informed consumer given the knowledge of g_t,g_{t-1},...and t_t,t_{t-1},...The same argument can be made for financial wealth (see appendix A).

assumption that these imperfectly informed consumers are aware of the intertemporal budget constraint of the government. We justify both assumptions by noting that consumers may value knowledge of the government budget constraint (e.g. because the level of government debt makes it worthwhile to 'pierce the government veil') without thinking it is necessary to know or to use the *exact* value of the aggregate variables in their calculations. Thus, we do not assume that consumers cannot obtain the aggregate information, we just assume that they do not value it enough to collect it or to use it (see Pischke, 1995 and Deaton, 1992)[5]. We capture this idea by assuming that consumers' expectations of their future tax liabilities are determined by their expectations of future (noisy) government expenditures. More precisely, we assume that $(E_{it}^{IM} - E_{it-1}^{IM})\left(\sum_{j=0}^{\infty}(1+r)^{-j}t_{it+j}\right) = (E_{it}^{IM} - E_{it-1}^{IM})\left(\sum_{j=0}^{\infty}(1+r)^{-j}g_{it+j}\right)$ where E_{it}^{IM} is the expectations operator conditional on the information set I_{it}^{IM} and E_{it-1}^{IM} is the expectations operator conditional on the information set I_{it-1}^{IM}.

Our assumptions have the following implications,

1) Since n is large and since there is independence across consumers, we have that the macro structure and the micro structure are compatible.

2) The change in consumption of each of the the n consumers can be written as (see appendix A),

[5] As noted by Deaton (1992-page 40), individuals have public access to aggregate variables. There is however no guarantee that they will use such information.

$$(9)\ \Delta c_{it} = r(1+r)^{-1}\sum_{j=0}^{\infty}(1+r)^{-j}(E_{it}^{k} - E_{it-1}^{k})(y_{it+j} - t_{it+j})$$

where i=1,…,n and E_{it}^{k} is the expectations operator conditional on the information set I_{it}^{k} (where k=PE or IM). Consumers make their decision at the end of t using the information set I_{it}^{k} available at the end of t.

3) To incorporate the assumption that consumers do not observe aggregate variables, we use Eq. (1), (2), (4), (5), (6) and (8) to derive the following equations (see appendix B),

$$(10)\quad \Delta y_{it} = y + \eta_{it} - \theta\eta_{it-1}$$

$$(11)\quad \Delta g_{it} = g + \mu_{it} - \phi\mu_{it-1}$$

where the parameters θ (>0) and φ (>0) are functions of the (relative) variances of aggregate (permanent and transitory) shocks and idiosyncratic shocks (see appendix B). η_{it} and μ_{it} reflect the innovations in consumer i's own pre-tax income and perception of government expenditures. These innovations are white noise and have a variance that is constant across consumers. In appendix B we show that the higher is the variance of the permanent shocks in pre-tax income and government expenditures, the lower will be the values of θ and φ. In this case, as we can see from (10) and (11), innovations in period t that lead to changes in y_{it} and g_{it} will almost not be undone in the next period.

Consumers thus experience shocks without knowing the exact nature of these innovations (aggregate persistent, aggregate temporary or idiosyncratic). Obviously, this is problematic when they calculate their permanent income. Since temporary aggregate and individual-specific shocks are white noise (Eqs. (1), (2), (6), (7) and (8)), they call for a small adjustment of permanent income. Aggregate variables may however also be affected by permanent shocks (see Eqs. (4) and (5)) which demand a larger adjustment of permanent income.

Perfectly informed consumers.

If consumers had perfect information about the aggregate economy and thus had information set I_{it}^{PE}, aggregate per capita consumption would be given by (see appendix C),

$$(12)\quad \Delta c_t = r(1+r)^{-1}(\varepsilon_t^y - \varepsilon_t^g) + v_t^y - v_t^g$$

This is the standard result that not only at the individual level (see Eq. 9) but also in the aggregate, perfectly informed permanent income consumers only respond to unanticipated shocks ('surprises'). Since they incorporate the government budget constraint, tax shocks do not enter Eq. (12).

Imperfectly informed consumers.

In appendix D we show that if all consumers are imperfectly informed but do take into account the government budget constraint, the change in per capita consumption is given by,

$$(13)\quad \Delta c_t = b^y a^y - b^g a^g + b^y \Delta y_t - b^g \Delta g_t + b^y \sum_{j=1}^{\infty} \theta^j \Delta y_{t-j} - b^g \sum_{j=1}^{\infty} \phi^j \Delta g_{t-j}$$

where $a^y = -(1-\theta)^{-1} y, a^g = -(1-\phi)^{-1} g$, $b^y = (1+r-\theta)(1+r)^{-1}$ and $b^g = (1+r-\phi)(1+r)^{-1}$.

As can be seen from (13), the lack of information on aggregate variables causes consumption to be excessively sensitive to (lags in) income (see Pischke 1995 and Demery and Duck 2000). Consumption is also excessively sensitive to (lags in) government expenditures. Consumption does not respond only to 'surprises' in income and government expenditures as is the case when consumers have complete information (see Eq. (12)). For instance, if in period t there is a permanent shock in government expenditures, this will lead to a change in g_{it}, g_{it+1},.... Consumers take this into account when determining consumption since they incorporate the government budget constraint. They will be uncertain about how much government expenditures have augmented since these expenditures are observed with noise. Consumers will therefore interpret the shock only partly as a permanent shock in aggregate g_t and will thus underestimate the persistence of the shock. They will change their estimate of permanent income insufficiently. Consumption in t will adjust, but the adjustment will be too small. The reaction depends on the relative variance of idiosyncratic and temporary aggregate versus permanent aggregate shocks. This can be seen from equation (13) where the change in consumption in period t due to a change in government expenditures in t depends on b^g. From appendix B we can derive that if the variance of v_t^g is small compared to the variances of ε_t^g and ε_{it}^g, the consumer will expect that shocks are mainly temporary or idiosyncratic and $\phi(>0)$ will be relatively large. At the aggregate level we then have $b^g = (1+r-\phi)(1+r)^{-1}$ and the reaction (of all consumers) to aggregate (permanent or temporary) shocks within the same period will be relatively small. In the next period(s), as the effect of a permanent shock is observed to persist, consumers will again be surprised and permanent income and consumption will be adjusted again. Consumption will then be excessively sensitive to lagged changes in government expenditures. The intuition for income is similar to that for government expenditures.

Note that our results on excess sensitivity of income and government expenditures are dependent on the imperfect information assumption at the individual level and on the aggregation bias that results from it. Consumption changes are unpredictable at the individual level, not at the aggregate level. We refer to appendix D for more on this. Working with a representative permanent income consumer would never generate excess sensitivity because the optimization problem of a representative consumer always implies unpredictable consumption changes at the aggregate level.

Note also that our results are not dependent on the presence of temporary aggregate shocks. We assume specifications like (1) and (2) with permanent and transitory shocks to leave room for predictability of income and government expenditure changes, given the random walk assumption of the permanent components. Assuming more general univariate processes for income and government expenditure changes would give similar results leading to private consumption changes that are excessively sensitive to (lagged) pre-tax income and government expenditure changes [6].

3. *Methodology and results.*

3.1 A testable consumption function, estimation issues and data issues.

Specifications in which the change in private consumption is a function of current and lagged changes in pre-tax income and government expenditures can be obtained from different

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[6] Demery and Duck (2000) obtain a similar result using an ARMA(p,q) for the change in aggregate income. In fact as long as Δy_{it} and Δg_{it} are covariance-stationary we can (by using the Wold theorem – see Hamilton 1994) write equations of the form $\Delta y_{it} = \mu + \psi(L)u_{it}$ where u_{it} is white noise and where $\psi(L)$ is an infinite-order lag polynomial with $\psi_0=1$ (and similar for Δg_{it}). Assuming the underlying parameters of the aggregate processes are such that the lag polynomials are invertible, we can obtain an expressions similar to (13).

assumed income and government expenditure processes. Thus, it may be too restrictive to test the model by imposing the assumed processes for income and government expenditures on the data [7]. In appendix E we therefore report the results of estimating Eq.(13) directly (i.e. the estimates for θ and ϕ) while in this section we estimate a more general consumption function,

$$(14)\quad \Delta c_t = \gamma + \delta_0^y \Delta y_t + \delta_0^g \Delta g_t + \sum_{j=1}^{q} \delta_j^y \Delta y_{t-j} + \sum_{j=1}^{q} \delta_j^g \Delta g_{t-j} + \varepsilon_t^c$$

where γ is a constant and the expected signs of the parameters are $\delta_0^y > 0$, $\delta_0^g < 0$, $\delta_j^y > 0$ (for j=1,...,q) $\delta_j^g < 0$ (for j=1,...,q). Note that strictly speaking an infinite number of lagged Δy_t and Δg_t terms should be included in (14), which is obviously not possible. Therefore, we follow the approach of Demery and Duck (2000) and we add a finite number q of lags. Given that enough lags of Δy_t and Δg_t are included we should find no increase in likelihood if lagged Δc_t terms are added to this equation. Lags of Δc_t added to our empirical specification (14) are indeed never significant and do not affect our results. All our results reported (for all countries) are for q=1 since using Wald tests we can easily reduce the number of lags to that number. We never find significant parameter estimates on Δy_{t-j} and Δg_{t-j} for j=2,3,.... [8]. Given their endogeneity we must instrument Δy_t and Δg_t. The error term ε_t^c can be interpreted as a preference shock or as resulting from transitory components in the level of consumption. In the latter case it will follow an MA(1) process and instruments must be

[7] Pischke finds that individual income changes are well described by an MA(2) Process. As noted by Deaton (1992), an MA(1) process for individual income changes, as in Eq. (10), may be a relatively good approximation to reality. This says nothing of Eq. (11) for government expenditures however

[8] Note that specification q=2 does generally not strongly affect the significance of our results (the standard errors and parameter estimates of the current value and first lag of Δy and Δg remain relatively stable). The significance of our results starts deteriorating however when more (insignificant) lags are added.

lagged at least twice to obtain consistent estimation. Stated differently, the variables Δy_{t-1} and Δg_{t-1} must also be instrumented. This is also necessary if we consider the possibility of time aggregation (Working 1960). We use lags of $\Delta c_t, \Delta y_t$ and Δg_t as instruments in our regressions.

To estimate (14) we use semi-annual data for 3 high-debt OECD countries (Belgium, Italy and Greece) from 1973:1 –1997:2. We take 1973 as a first observation because it marks the first oil shock. This crisis and the fiscal and debt problems that it provoked, may have increased consumers' awareness of government issues. Using semi-annual data gives the possibility to estimate (14) over relatively small subperiods and to capture the dynamics implied by the lag structure in (14). We would prefer quarterly data but these are not available for all countries for all necessary variables. More details on the construction of the used variables are given in table 1.

Table 1. Data and data sources.

c_t	Real aggregate per capita private consumption. Taken from OECD Economic Outlook (CD ROM 2001 Vol.2) and reported in real terms with code CPV (deflated by implicit deflator for aggregate consumption, 1995=100).
y_t	Real per capita pre-tax income. Calculated using variables from OECD Economic Outlook (CD ROM 2001 Vol.2). Multiply the wage rate in the business sector (with code WR) by employment in the business sector (with code ETB) and then add government wages (with code CGAA). Then deflate this sum by the private consumption deflator (with code PCP, 1995=100). Note that this measure cannot be constructed on a semi-annual basis for Belgium before 1980. For Belgium we therefore use wages and salaries (code WAGE in OECD statistics) which is basically the same and which is available from 1973 onwards. We also deflate this series with PCP. Estimations with both measures (after 1980) are practically identical. We name these proxies Y2. Since this type of proxy cannot be constructed for Greece we also proxy pre-tax income by real per capita GDP with code GDP (see for instance Evans and Karras 1996, 1998). We name this proxy Y1.
g_t	Real per capita government consumption plus real per capita government investment. Taken from OECD Economic Outlook (CD ROM 2001 Vol.2). Real government consumption is reported with code CGV (deflated by implicit deflator for government consumption, 1995=100). Real government investment is reported with code IGV (deflated by implicit deflator for government investment, 1995=100).

Note: all variables are seasonally adjusted. Per capita measures are obtained after dividing by total population. Semi-annual data for total population are constructed by dividing the data for the population between 15 and 64 years of age (which are available on a semi-annual basis in OECD Economic Outlook CD ROM 2001 Vol.2 with code POPT) by the ratio of the population of 15-64 to total population, which can be calculated on a yearly basis. The latter is calculated using total population (available on a yearly basis only) from OECD Economic Outlook CD ROM 2001 vol.2 (with code POP).

3.2 Estimation results.

Note first that we do reject the hypothesis of a unit root in all series used for Δc_t, Δy_t and Δg_t for all countries and over all sample periods considered. Unit root tests are not reported but they are available upon request.

We report our results for Eq. (14) with q=1 using the instrumental variables (IV) approach in tables 2 (Italy), 3 (Greece) and 4 (Belgium). Note that specifications with q=2 do generally lead to the same conclusions, though the additional lags are never significant. For each country we consider subperiods and a number of different variables to proxy pre-tax income (Y1,Y2). We use instrument sets containing the second to the fifth lag of Δc_t, Δy_t and Δg_t when estimating our equation. Note that changing the number of instruments in all these cases does only marginally affect our results.

Table 2. IV estimates of (14) with q=1 for Italy.

Coefficient	(1) 1973:01-1997:02 with Y1 [b]	(2) 1973:01-1997:02 with Y2 [b]	(3) 1982:01-1997:02 with Y2 [b]
δ_0^y	0.219** (0.082)	0.479* (0.295)	0.493** (0.169)
δ_0^g	-0.829** (0.239)	-0.716** (0.248)	-0.440* (0.261)
δ_1^y	0.207** (0.095)	0.358** (0.183)	0.402* (0.205)
δ_1^g	0.384* (0.220)	0.156 (0.345)	0.133 (0.248)
N obs.	50	50	32
R^2	0.371	0.124	0.455
DW[a]	2.270	1.678	1.561

Notes: Newey-West standard errors between brackets (lag truncation = 3). * indicates significance at the 10% level. ** indicates significance at the 5% level. The instrument set contains lags 2 to 5 of Δc_t, Δy_t and Δg_t.
[a] DW indicates the Durbin-Watson test statistic. [b] Y1 is real per capita GDP used as a proxy for y_t, Y2 is real per capita total wages used as a proxy for y_t.

Table 2 presents the results for Italy. In the first equation we estimate (14) over the period 1973:01-1997:02 with real per capita GDP (Y1) as a proxy for y_t. All coefficients have the expected sign, except δ_1^g which is positive. This contradicts the intuition of the model. In the second equation where we use real per capita total wages (Y2) as a proxy for y_t, our coefficient estimates are considerably different. δ_1^g is now insignificant.

Note that if we estimate (14) over smaller sample periods as in the third equation, the results remain the same. We report the results for the period 1982:01-1997:02 because in 1982 the net debt to GDP ratio reached 50% and a long uninterrupted increase followed (see figure 1 in section 1). The results are robust to estimation over other sample periods starting in the early to mid eighties however.

We thus find excess sensitivity with respect to current pre-tax income and government expenditures and with respect to the first lag in income. Lagged changes in government expenditures do not enter the regression in a significant way. A possible explanation for this could be that most of the adjustment in government expenditures occurs immediately, for instance in the first quarter after the shock. This could also explain why δ_0^g is quite large in absolute terms (especially in Eqs (1) and (2)). Consumers may think that changes in government expenditures are usually rather persistent so that adjustment occurs fast and over a small period. Quarterly data might be more useful to capture the adjustment. However, due to data limitations, only semi-annual data are available.

Note that the usual interpretation for the observation that private consumption responds negatively to government expenditure changes would be that private consumption and government expenditures are Edgeworth substitutes. As noted earlier, at the aggregate level, there may be reason to question such a relationship. The finding that lagged pre-tax income enters the regression significantly provides support for the alternative explanation considered in this paper.

Additional support for the idea of this paper is given by the results for Greece in table 3. For Greece we can only use Y1 as a proxy for real per capita pre-tax labour income. In the first equation we estimate (14) over the period 1973:01-1997:02. No variables are significant except current income. The Durbin Watson statistic being smaller than 1, our consumption function seems to be rather misspecified for this period. Note however that during the seventies and most of the early eighties the debt ratio in Greece remained relatively low (see figure 2 in section 1). If we estimate (14) over the period 1985:01-1997:02, a period of high debt levels (almost exceeding 50% in 1985) and a very strong increase in the debt ratio, another picture emerges. We find results strongly in line with these for Italy. Again, we find excess sensitivity with respect to current pre-tax income and government expenditures and the first lag in income. We do not find a significant influence from lagged changes in government expenditures. The Durbin-Watson statistic is much closer to 2 this time.

Table 3. IV estimates of (14) with q=1 for Greece.

Coefficient	(1) 1973:01-1997:02 with Y1[b]	(2) 1985:01-1997:02 with Y1[b]
δ_0^y	0.235** (0.075)	0.358** (0.086)
δ_0^g	-0.191 (0.353)	-0.822** (0.343)
δ_1^y	0.011 (0.047)	0.331** (0.103)
δ_1^g	0.255 (0.206)	0.001 (0.191)
N obs.	50	26
R^2	0.583	0.670
DW[a]	0.966	1.605

Notes: Newey-West standard errors between brackets (lag truncation = 3). * indicates significance at the 10% level. ** indicates significance at the 5% level. The instrument set contains lags 2 to 5 of Δc_t, Δy_t and Δg_t.
[a] DW indicates the Durbin-Watson test statistic. [b] Y1 is real per capita GDP used as a proxy for y_t.

Far less convincing are the results for Belgium, however (table 4). From figure 1 in section 1 we note that of all countries considered Belgium has the highest debt level. Our results are not reflecting this. In the first equation (full sample with Y1) only current income enters significantly. The Durbin-Watson test is very low suggesting model misspecification. In the second equation we use real per capita total wages (Y2) as a proxy for y_t. Again only current income is significant, though current government expenditures have the expected sign. Using subsamples does not improve the results.

Table 4. IV estimates of (14) with q=1 for Belgium.

Coefficient	(1) 1973:01-1997:02 with Y1 [b]	(2) 1973:01-1997:02 with Y2 [b]	(3) 1985:01-1997:02 with Y2 [b]
δ_0^y	0.313** (0.074)	0.434** (0.136)	0.589** (0.280)
δ_0^g	0.052 (0.116)	-0.156 (0.108)	0.336 (0.559)
δ_1^y	0.015 (0.047)	0.008 (0.114)	-0.204 (0.148)
δ_1^g	0.053 (0.082)	0.002 (0.081)	-0.503 (0.295)
N obs.	50	50	26
R^2	0.662	0.464	0.275
DW[a]	1.096	1.075	1.003

Notes: Newey-West standard errors between brackets (lag truncation = 3). * indicates significance at the 10% level. ** indicates significance at the 5% level. The instrument set contains lags 2 to 5 of Δc_t, Δy_t and Δg_t.
[a] DW indicates the Durbin-Watson test statistic. [b] Y1 is real per capita GDP used as a proxy for y_t, Y2 is real per capita total wages used as a proxy for y_t.

Summarizing, our results are far from conclusive. There are indications though that our model can explain certain observations of excess sensitivity of consumption to income and government expenditures in some countries during some periods.

4. Conclusions.

In this paper we consider an alternative to the Edgeworth interpretation of excess sensitivity of aggregate consumption to anticipated changes in government expenditures building on the idea of imperfect information as developed by Pischke (1995).

We consider a model where rational utility maximizing permanent income consumers are assumed to be imperfecly informed about the aggregate economy. They do however incorporate the intertemporal budget constraint of the government. At the aggregate level this generates excess sensitivity of private consumption changes to predictable pre-tax income and government expenditure changes.

In the empirical section we estimate a consumption function implied by the model for three high debt countries (Italy, Greece and Belgium). We concentrate on high debt countries since it can be expected that consumers in these countries may value information on government budget issues more than consumers in other countries.

Our findings suggest that in Italy and Greece (especially during periods of high debt accumulation) private consumption changes are indeed excessively sensitive to (lagged) changes in pre-tax income and government expenditures. Our results are not fully conclusive however since for Belgium there are no indications that the model provides a satisfying approximation to reality.

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Appendix A: Derivation of (9).

Given the quadratic utility assumption and the assumption that the subjective rate of time preference equals the constant real interest rate, the first order condition at time t can be written as (∀j),

$$\text{(A1)} \quad E_{it}^{k} c_{it+j} = c_{it}$$

with k=IM (or PE if consumers are perfectly informed) and i=1,...,n. This is the standard random walk result (see Hall 1978). The period t budget constraint of these consumers can be written as,

$$\text{(A2)} \quad c_{it} + w_{it}(1+r)^{-1} = y_{it} - t_{it} + w_{it-1}$$

where w_{it} is consumer i's financial wealth (including government bonds) measured at the end of period t. Solving (A2) forwards, imposing a solvency condition and taking expectations E_{it}^k leads to,

$$\text{(A3)} \quad \sum_{j=0}^{\infty}(1+r)^{-j} E_{it}^k c_{it+j} = w_{it-1} + \sum_{j=0}^{\infty}(1+r)^{-j} E_{it}^k (y_{it+j} - t_{it+j})$$

Substituting (A1) into (A3) we obtain,

$$\text{(A4)} \quad c_{it} = r(1+r)^{-1}\left(\sum_{j=0}^{\infty}(1+r)^{-j} E_{it}^k (y_{it+j} - t_{it+j}) + w_{it-1}\right)$$

Substituting the budget constraint (A2) written for t-1 into (A4), we obtain

$$\text{(A5)} \quad c_{it} = r(1+r)^{-1}\sum_{j=0}^{\infty}(1+r)^{-j} E_{it}^k (y_{it+j} - t_{it+j}) + r(y_{it-1} + w_{it-2} - t_{it-1} - c_{it-1})$$

Lagging (A4) one period, multiplying both sides by 1+r and extracting the term for j=0 from the summation, we obtain,

$$\text{(A6)}\quad (1+r)c_{it-1} = rw_{it-2} + r(y_{it-1} - t_{it-1}) + r\sum_{j=1}^{\infty}(1+r)^{-j}E_{it-1}^{k}(y_{it+j-1} - t_{it+j-1})$$

Rearranging terms, (A6) can also be written as,

$$\text{(A7)}\quad c_{it-1} = -rc_{it-1} + rw_{it-2} + r(y_{it-1} - t_{it-1}) + r(1+r)^{-1}\sum_{j=0}^{\infty}(1+r)^{-j}E_{it-1}^{k}(y_{it+j} - t_{it+j})$$

Subtracting (A7) from (A5) we obtain (9) in the main text.

Appendix B: Derivation of (10) and (11).

We focus on the derivation of (10), the derivation of (11) is completely identical. Substituting (1) and then (4) into (6), and rewriting the result in first differences, we obtain,

$$\text{(B1)}\quad \Delta y_{it} = y + v_t^y + \varepsilon_t^y - \varepsilon_{t-1}^y + \varepsilon_{it}^y - \varepsilon_{it-1}^y$$

In (B1) we have a combination of various white noise errors. To capture the fact that consumers do not observe aggregate variables, we can write this combination of errors as an MA(1) process,

$$\text{(B2)}\quad v_t^y + \varepsilon_t^y - \varepsilon_{t-1}^y + \varepsilon_{it}^y - \varepsilon_{it-1}^y = \eta_{it} - \theta\eta_{it-1}$$

with η_{it} being a white noise term with a variance that is constant across consumers. The value of θ that ensures the white noise structure of η_{it}, is obtained if we equate the first-order autocorrelations of both sides of (B2) [9]. This leads to

$$\text{(B3)}\quad \theta^2+1=\left(\frac{\sigma^2_{v^y}+2\sigma^2_{\varepsilon^y}+2\sigma^2_{\varepsilon^y_i}}{\sigma^2_{\varepsilon^y}+\sigma^2_{\varepsilon^y_i}}\right)\theta$$

where $\sigma^2_{\varepsilon^y_i}$, $\sigma^2_{\varepsilon^y}$ and $\sigma^2_{v^y}$ are the unconditional variances of ε^y_{it}, ε^y_t and v^y_t respectively.

Eq. (B3) is a quadratic of the form $a\theta^2+b\theta+c=0$ with a=c=1 and $b=-\left(\frac{\sigma^2_{v^y}+2\sigma^2_{\varepsilon^y}+2\sigma^2_{\varepsilon^y_i}}{\sigma^2_{\varepsilon^y}+\sigma^2_{\varepsilon^y_i}}\right)$. To have at least one real solution θ* the condition must hold that $b^2\geq 4$. This condition always holds. Note that from (B3) it is obvious that θ>0. The two roots are given by $\theta_1=(1/2)(-b+\sqrt{b^2-4})$ and $\theta_2=(1/2)(-b-\sqrt{b^2-4})$. Using the solution θ* into (B2) and substituting this into (B1) we obtain,

$$\text{(B4)}\quad \Delta y_{it}=y+\eta_{it}-\theta^*\eta_{it-1}$$

which equals (10) in the main text. Note that only with the root θ_2 we have $\theta_2<1$. This is the invertibility condition that guarantees that from (B4) we can write,

[9] Where we use the assumption that there is no correlation between permanent aggregate, temporary aggregate and idiosyncratic shocks (not contemporaneous and not at any lead nor lag).

(B5) $\eta_{it} = a^y + \sum_{j=0}^{\infty} \theta^j \Delta y_{it-j}$

where $\theta = \theta^* = \theta_2$ and $a^y = -(1-\theta)^{-1} y$.

In exactly the same way we can derive Eq. (11) in the main text. We can write,

(B6) $v_t^g + \varepsilon_t^g - \varepsilon_{t-1}^g + \varepsilon_{it}^g - \varepsilon_{it-1}^g = \mu_{it} - \phi\mu_{it-1}$

with $\phi>0$. Note that only the second root ϕ_2 will satisfy the invertibility condition $\phi_2<1$ so that we can write from (11),

(B6) $\mu_{it} = a^g + \sum_{j=0}^{\infty} \phi^j \Delta g_{it-j}$

where $\phi = \phi^* = \phi_2$ and $a^g = -(1-\phi)^{-1} g$.

Appendix C: Derivation of (12).

From the main text we know that perfectly informed consumers observe all aggregate variables (information set I_{it}^{PE}). We start from Eq. (9) into which we substitute Eqs. (6) and (7) written for period t+j. Substituting out taxes by means of the budget constraint (3), we obtain, (C1) $\Delta c_{it} = r(1+r)^{-1}\sum_{j=0}^{\infty}(1+r)^{-j}(E_{it}^{PE} - E_{it-1}^{PE})\left[y_{t+j} - g_{t+j} + \varepsilon_{it+j}^{y} - \varepsilon_{it+j}^{t}\right]$

Using Eqs. (1),(2),(4) and (5) into (C1) we obtain,

(C2) $\Delta c_{it} = r(1+r)^{-1}(\varepsilon_{t}^{y} - \varepsilon_{t}^{g} + \varepsilon_{it}^{y} - \varepsilon_{it}^{t}) + v_{t}^{y} - v_{t}^{g}$

Summing Eq. (C2) over all n consumers makes the idiosyncratic shocks disappear. We obtain Eq. (12) in the main text.

Appendix D: Derivation of (13).

At the end of period t imperfectly informed consumers have information set I_{it}^{IM}. We also know that these consumers form expectations according to $(E_{it}^{IM} - E_{it-1}^{IM})\left(\sum_{j=0}^{\infty}(1+r)^{-j}t_{it+j}\right) = (E_{it}^{IM} - E_{it-1}^{IM})\left(\sum_{j=0}^{\infty}(1+r)^{-j}g_{it+j}\right)$. Substituting this into Eq. (9) with k=IM we obtain,

(D1) $\Delta c_{it} = r(1+r)^{-1}\sum_{j=0}^{\infty}(1+r)^{-j}(E_{it}^{IM} - E_{it-1}^{IM})(y_{it+j} - g_{it+j})$

We first put (D1) into a different form. Note that we can write (D1) as,

(D2)

$$\Delta c_{it} = \sum_{j=0}^{\infty}(1+r)^{-j}(E_{it}^{IM} - E_{it-1}^{IM})(y_{it+j} - g_{it+j}) - (1+r)^{-1}\sum_{j=0}^{\infty}(1+r)^{-j}(E_{it}^{IM} - E_{it-1}^{IM})(y_{it+j} - g_{it+j})$$

The second term of (D2) can also be written as $-\sum_{j=0}^{\infty}(1+r)^{-j}(E_{it}^{IM}-E_{it-1}^{IM})(y_{it+j-1}-g_{it+j-1})$.

Plugging this into (D2) we obtain,

$$\text{(D3)}\quad \Delta c_{it}=\sum_{j=0}^{\infty}(1+r)^{-j}(E_{it}^{IM}-E_{it-1}^{IM})(\Delta y_{it+j}-\Delta g_{it+j})$$

Substituting Eqs. (10) and (11) into (D3) leads to

$$\text{(D4)}\quad \Delta c_{it}=b^{y}\eta_{it}-b^{g}\mu_{it}$$

where $b^{y}=(1+r-\theta)(1+r)^{-1}$ >0 (given that it follows from app. B that 0<θ<1), $b^{g}=(1+r-\phi)(1+r)^{-1}$>0 (given that it follows from app.B that 0<ϕ<1).

Substituting Eqs. (B5) and (B6) into (D4) and aggregating over the n consumers leads to (13) in the main text.

Note from (D4) that at the individual level the change in consumption is white noise. If in period t a persistent aggregate shock occurs, consumption will be adjusted. In the following period, due to the persistence of the shock, the consumer will again be surprised. The consumption change is thus orthogonal to the information set of the imperfectly informed consumer, but not orthogonal to last period's aggregate shock. Note that it is assumed that consumers never learn about the aggregate persistent shocks. Given that calculations by Pischke (1995) suggest that for consumers in the US the costs of obtaining aggregate information may be far greater than the benefits, it seems plausible that consumers choose to remain ignorant indefinitely (see Deaton 1992).

At the aggregate level consumption changes are not white noise, however. If we aggregate Eq. (D4) over all n consumers we obtain $\Delta c_t = b^y \eta_t - b^g \mu_t$ where $\eta_t = (1/n)\sum_{i=1}^{n} \eta_{it}$ and $\mu_t = (1/n)\sum_{i=1}^{n} \mu_{it}$. Note now that while η_{it} and μ_{it} are white noise, η_t and μ_t are not. To see this, we can use (B2) and (B6). The presence of the terms in ε_{it}^y and ε_{it}^g forces the selection of particular values for θ and ϕ to guarantee white noise in each η_{it} and μ_{it}. On aggregation, since the terms in ε_{it}^y and ε_{it}^g disappear and since θ and ϕ are the same, η_t and μ_t will not be white noise.

Appendix E: Estimation of Eq. (13).

In this appendix we estimate the parameters θ and ϕ directly from Eq. (13). We first rewrite Eq. (13) in a different form. Note that from (B4) we can write $\eta_{it} = (\Delta y_{it} - y)(1-\theta L)^{-1}$ where L is the lag operator. We can do the same for government expenditures, so that we have $\mu_{it} = (\Delta g_{it} - g)(1-\phi L)^{-1}$. Substituting both expressions in (D4), multiplying both sides by $(1-\theta L)(1-\phi L)$ and aggregating[10], we obtain,

$$\text{(E1)}\quad \begin{aligned}\Delta c_t &= (\theta+\phi)\Delta c_{t-1} - \theta\phi\Delta c_{t-2} + (1-\theta)(1+r-\phi)(1+r)^{-1} g - (1-\phi)(1+r-\theta)(1+r)^{-1} y \\ &+ (1+r-\theta)(1+r)^{-1}((1-\phi L)\Delta y_t) - (1+r-\phi)(1+r)^{-1}((1-\theta L)\Delta g_t)\end{aligned}$$

In table E1 we report the results of the estimation of Eq.(E1) over the indicated sample periods for Italy, Greece and Belgium. We use real per capita wages for y_t for Italy and Belgium (Y2) and real per capita GDP for Greece (Y1). We report the results obtained for the estimation of θ and ϕ with imputed values for y, g (which are the averages of Δy_t and Δg_t over the sample period) and r. We set r=0.02 in all estimations, but the results are not sensitive to the use of other values for r in the interval 0.01-0.05.

Table E1. IV estimates of Eq. (E1) with r=0.02.

	Coefficient		N Obs
	θ	ϕ	
(1)	0.747**	0.158	50
Italy (1973:01-1997:02)	(0.102)	(0.239)	
(2)	0.813**	0.166	32
Italy (1982:01-1997:02)	(0.170)	(0.252)	
(3)	0.714**	0.694**	50
Greece (1973:01-1997:02)	(0.073)	(0.136)	
(4)	0.567**	0.447**	26
Greece (1985:01-1997:02)	(0.072)	(0.149)	
(5)	0.732**	0.775**	50
Belgium (1973:01–1997:02)	(0.165)	(0.110)	
(6)	0.637**	0.996**	26
Belgium (1985:01-1997:02)	(0.174)	(0.228)	

Notes: Newey-West standard errors between brackets (lag truncation = 3). * indicates significance at the 10% level. ** indicates significance at the 5% level. The instrument set contains lags 2 to 5 of Δc_t, Δy_t and Δg_t. Y2 is used for Italy and Belgium; Y1 is used for Greece (see table 1 for exact definitions).

From the results in table E1 we note that the estimates for θ are significant and plausible (i.e. between 0 and 1) in all cases considered. The estimates for ϕ are insignificant for Italy but have significant values between 0 and 1 for the other countries. Note that whereas the point estimates for θ are very similar over the different countries, the point estimates for ϕ are very different (close to 0 for Italy, close to 0.5 for Greece and close to 1 for Belgium). The point estimates for θ are all relatively close to 1. In terms of the model this implies a relatively high variance of the transitory and/or individual-specific components in income (see Eq. 13) so

[10] Note that the conditions $\theta<1$ and $\phi<1$ are not necessary here (values larger than 1 would be rather unrealistic however). This also implies that b^y and b^g could be negative.

that consumers may think that changes in income are usually not very persistent. Permanent income and consumption will not be adjusted very much in the period of the shock and will adjust relatively slowly in the consequent periods.

The result that ϕ is relatively low in Italy is more or less in accordance with the results found in table 2 where a more general consumption function is estimated (Eq. 14) and where the lagged changes in government expenditures are insignificant. The variance of permanent shocks in government expenditures may be relatively high so that consumers may think that changes in government expenditures are usually rather persistent so that adjustment occurs fast and over a small period (see Eq. 13 in the paper to see this). More or less the same conclusion can be drawn for Greece if we estimate over the period 1985:01-1997:02 (the point estimate for ϕ is still lower than 0.5 and significant in this case). In the other cases (Greece 1973:01-1997:02 and Belgium over both sample periods considered) we find relatively high values for ϕ. These cases coincide with general estimations (Eq.14) of very poor quality (see the discussion in the main text). Rather than taking these estimates for ϕ (and the large differences across countries which are implausible) seriously, we conclude that these results may also be due to the fact that the estimation of (13) and (E1) may be too restrictive (especially the process for government expenditures assumed in the model – see footnote 7).

Government debt and the excess sensitivity of private consumption to current income: an empirical analysis for OECD countries [(*)]

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Abstract.

Empirical studies typically find that private consumption is much more sensitive to changes in current disposable income than is predicted by Hall's (JPE, 1978) permanent income hypothesis. Standard explanations for this "excess sensitivity" of private consumption refer to liquidity constraints and/or myopia. Elaborating on existing theoretical literature, which suggests that the incidence of liquidity constraints and the degree of myopia may be affected by the government debt ratio, this paper investigates the role of government debt in the degree of excess sensitivity. Using a panel of OECD countries in the 1990s, we estimate a consumption function with the degree of excess sensitivity depending on the government debt ratio and the degree of financial liberalisation. We find that a higher government debt leads to more excess sensitivity. This result supports the idea that a higher debt induces private lenders to tighten credit conditions, which raises the incidence of liquidity constraints. As to the effects of financial liberalisation on excess sensitivity in the 1990s, we obtain no clear evidence.

JEL Classifications: E21, E62, C33.

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1. Introduction.

Assuming rational forward-looking consumers and perfect capital markets, Hall (1978) has demonstrated that under the permanent income hypothesis consumption should follow a random walk. Most studies during the past two decades have, however, rejected this prediction. In particular, they have concluded that private consumption is more sensitive to current disposable income than is consistent with the permanent income hypothesis. In the literature several explanations for this "excess sensitivity" have been put forward. Typically these explanations come down to dropping one or more of Hall's assumptions. For example, some authors have referred to myopic behaviour from a significant part of the consumers, i.e. a deviation from the basic postulate of rational forward-looking agents (e.g. Flavin, 1985; Romer, 2001). Many others have attributed excess sensitivity to credit market imperfections and liquidity constraints, preventing rational consumers from realising their desired consumption (see e.g. the seminal work by Flavin, 1981 and 1985 and Campbell and Mankiw, 1990). Other potential explanations for observed excess sensitivity to income relate to precautionary savings (Carroll, 1992), imperfect information (Goodfriend, 1992; Pischke, 1995) and misspecification of the estimated consumption function (Campbell and Mankiw, 1990). The focus of this paper is on myopia and liquidity constraints.

Building on the idea of liquidity constraints, several authors have more recently endogenised the degree of "excess sensitivity". *Cross-sectionally*, Jappelli and Pagano (1989) and Campbell and Mankiw (1991) find that countries with better developed capital markets and easier access to credit have lower excess sensitivity of private consumption. Haliassos and Christou (2000) cannot reject that countries with high concentration and low efficiency in the banking sector have higher excess sensitivity. Evans and Karras (1998) show for 66 countries that the excess sensitivity of consumption to disposable income is lower in countries with high savings rates. A higher savings rate implies that consumers accumulate more wealth, which makes them less vulnerable to liquidity constraints. A number of papers have investigated the hypothesis that the deregulation of credit markets in many countries during the last decades has *over time* lowered the fraction of credit constrained consumers and the excess sensitivity of private consumption. Bayoumi and Koujianou (1990), Blundell-Wignall et al. (1995), McKiernan (1996) and Girardin et al. (2000) can confirm this hypothesis for several OECD countries (e.g. US, France). However, Campbell and Mankiw (1991) cannot. Finally, Bacchetta and Gerlach (1997) have demonstrated the role of (endogenous) liquidity constraints for private consumption from a different perspective. They show excess sensitivity of consumption to credit aggregates in the US, Canada, the UK, France and Japan. As to the evolution of excess sensitivity *over time*, they only observe a clear tendency of decline in the US. Despite financial liberalisation, they do not observe this tendency in the other countries.

This paper focuses on the role of government debt for the excess sensitivity of private consumption. To the best of our knowledge, the empirical literature on excess sensitivity has until now disregarded government debt. Theoretically, however, a number of obvious channels have been suggested. On the one hand, a (very) high or (rapidly) increasing government debt ratio may alert unconscious citizens and, as a consequence, reduce the fraction of myopic consumers. Excess sensitivity of private consumption to current income should then fall. On the other hand, high or rising government debt ratios imply an increase in households' future liabilities. Banks or other lenders may then reduce the amounts they lend, thereby raising the incidence of liquidity constraints and excess sensitivity of private consumption. In section 2 we develop these theoretical channels somewhat further and put them into a workable econometric framework. In particular, we derive an equation for the change in private consumption, with the degree of excess sensitivity being a function of the government debt ratio (and the degree of financial liberalisation). In section 3 we estimate this consumption equation for 15 to 19 OECD countries. We make use of panel estimation methods that allow us to correct for simultaneity and heterogeneity. Our results support the idea that a higher government debt ratio implies tighter credit conditions and an increase of excess sensitivity. This unfavourable effect from increasing government debt on credit conditions seems to exist especially when government debt is already at a high level. Section 4 summarises our main results and their implications.

2. *Government debt and the excess sensitivity of private consumption: a theoretical framework.*

Like many before, we take Campbell and Mankiw's (1990) methodology as our starting point. Campbell and Mankiw have extended Hall's (1978) Euler equation approach to consumption by allowing for two groups of consumers. One group consists of forward-looking permanent income consumers, the other consists of rational liquidity constrained consumers. In this paper we also take into account the possibility of myopic consumers.

The standard model of consumption behaviour considers the optimal consumption path of a representative rational consumer who can lend and borrow freely. Assuming that the real interest rate is constant and equal to the subjective rate of time preference, one obtains that in the optimum

$$(1)\quad U'(C_{t-1}) = E_{t-1}U'(C_t)$$

where C_t is the level of consumption in t, $U'(C_t)$ the marginal utility of consumption and E_{t-1} the expectations operator, conditional on information available at time $t-1$. If the marginal utility of consumption is linear, the change in consumption is unpredictable. Alternatively, if the utility function is iso-elastic and the aggregate consumption level is distributed log-normally, the growth rate of consumption cannot be forecasted (see also Girardin et al., 2000). One then obtains that

(2) $\Delta c_t = \alpha + \varepsilon_t$

where $\Delta c_t = lnC_t - lnC_{t-1}$ and ε_t is uncorrelated with lagged variables.

A large number of empirical studies have however rejected equation (2)[1]. Typically, the evidence supports an alternative specification in which consumption displays "excess sensitivity" to disposable income, that is

(3) $\Delta c_t = \alpha + \lambda_t \Delta y_t + \varepsilon_t$

with Δy_t the change in the log of current disposable income and λ_t the "excess sensitivity" parameter. As we have mentioned before, standard explanations for excess sensitivity refer to the existence of liquidity constrained and myopic consumers. The larger the fractions of these two groups, the higher λ will be. Furthermore, building on the idea of liquidity constrained consumers, several authors have emphasised the possibility of a time-varying λ due to financial deregulation in many countries, especially in the OECD (see the subscript t for λ in equation 3). In this paper we also endogenise λ, putting the role of government debt at the centre. Equation (4) summarises. For a specific functional form to be estimated, we refer to section 3.

$$(4)\ \lambda_t = \lambda(\overset{?}{b_t}, \overset{-}{FL_t})$$

In this equation FL_t stands for the degree of financial deregulation and b_t for the government debt to GDP ratio, both measured at the beginning of t. The sign above a variable indicates the expected effect of increases in that variabele on λ_t. Equation (4) first reflects the hypothesis that financial liberalisation reduces λ because it implies a smaller fraction of liquidity constrained consumers. As we have described in section 1, a majority of studies confirm this hypothesis. The effect of changes in the government debt to GDP ratio is theoretically ambiguous. On the one hand, a higher government debt may raise the fraction of liquidity constrained consumers, on the other hand it may reduce the fraction of myopic consumers. Underlying the former effect is the simple idea that higher government debt ratios imply an increase in households' future liabilities. Banks or other lenders may then reduce the amounts they lend, thereby raising the incidence of liquidity constraints (Hayashi, 1987; Yotsuzuka, 1987; Romer, 2001). Figure 1 provides some preliminary support for this idea. It shows that in the 1970s, for which data are available, the loan-to-value ratio for home purchases was typically higher in countries with a lower net government debt ratio[2].

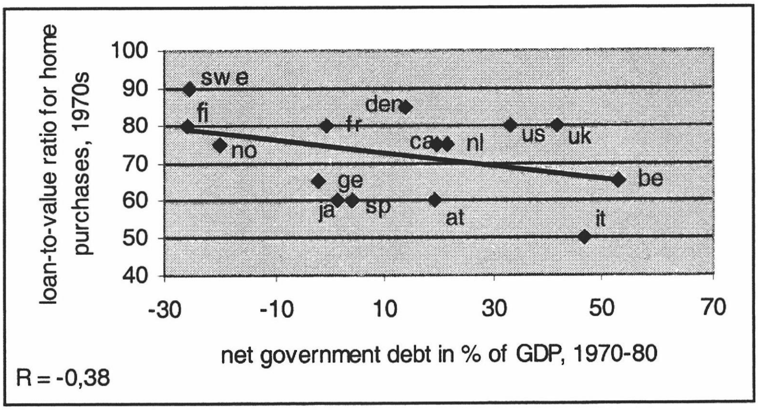
A higher government debt ratio may on the other hand reduce the fraction of myopic

[1] For references, see section 1.

[2] Romer (2001, p. 540) however raises doubts about the proposition that a high government debt induces tighter credit conditions. In his view it only arises when taxes are lump-sum. In the more realistic case of income taxes, borrowers will typically have to pay less taxes when they face difficulties to repay their (bank) loans, i.e. in bad times. Banks consequently know that their borrowers' share in repaying the government debt will be low, precisely at times when they may face trouble repaying their bank loans. Bond issues by the government are therefore likely to have only a small effect on borrowers' probabilities to repay private loans, and hence only a small effect on the amount that they can borrow.

consumers (Dalamagas, 1993a, 1994). The intuition is simple. To the extent that a very high or exploding government debt ratio gets more attention in political debate and/or the media, it may raise consumers' awareness of the future (future taxes). As a consequence, the share of myopic consumers will fall. So will excess sensitivity. This relationship is consistent with the results of existing empirical work that at high or exploding debt levels consumption behaviour will be more Ricardian (e.g. Nicoletti, 1989; Nicoletti, 1992; Dalamagas, 1993b, 1994; Slate et

Figure 1. Net government debt and loan-to-value ratio in 15 OECD countries in the 1970s.



Source: Loan-to-value ratio: Japelli and Pagano (1994); net government debt: OECD (2001). Included countries are Austria, Canada, Belgium, Denmark, Finland, France, Germany, Italy, Japan, The Netherlands, Norway, Spain, Sweden, UK and US. For four countries the data series for net government debt in 1970-1980 is incomplete: Austria (1980 only), Denmark (1980 only), Norway (1979-1980) and Spain (1976-1980). If we drop these four countries, correlation in figure 1 becomes –0.46.

al.,1995)[3]. Nicoletti (1988) has estimated private consumption functions for eight OECD countries over the 1961-85 period.

[3] Giavazzi et al. (2000) and Heylen and Everaert (2000), however, provide evidence that is inconsistent with this hypothesis. Furthermore, note that the results of Nicoletti (1988) and others do not prove that high government debt ratios raise consumer awareness of the future effects of fiscal policy. These results can also be derived from other hypotheses. For example, it may be that consumers are perfectly aware of the government budget constraint and the future tax implications of debt accumulation. But they may discount these future taxes only when the debt rate is very high. Only at a high debt rate they may rationally feel that the "day of reckoning" will still arrive during their lives.

He finds that expected future taxes are discounted much more strongly in consumer behaviour in highly indebted countries (Belgium, Italy) than in countries where the fiscal stance is sustainable. Whereas the traditional Keynesian view seems to be appropriate in low debt countries, there is some support for the Ricardian view on consumption in high debt countries. Nicoletti (1992) shows for Belgium that tax discounting is time-varying and increasing with the debt ratio. Dalamagas (1993b, 1994) provides evidence that in low debt countries consumers respond to a reduction in the ratio of taxes to the government deficit by increasing consumption. In high debt countries they don't. They may even reduce consumption. Finally, Slate et al. (1995) have carried out a number of experiments to test Ricardian equivalence under uncertainty. The results of these experiments suggest that the response of people to fiscal deficits tends to be Keynesian when the probability of debt repayment is low. If, on the other hand, the probability of debt repayment is high, people act much more in a Ricardian way.

3. Government debt, financial liberalisation and the excess sensitivity of private consumption: an empirical analysis.

3.1. Basic set-up.

In this section we test the model described by equations (3) and (4) using panel data for 15 to 19 OECD countries in the period 1990-99. Equation (5) reflects this panel data set-up. Equation (6) describes a specific functional form for λ_t. We adopt a straightforward linear specification.

$$(5)\quad \Delta c_{jt} = \alpha_j + \lambda_{jt}\Delta y_{jt} + \varepsilon_{jt}$$

$$(6)\quad \lambda_{jt} = \beta_0 + \beta_1 b_{jt} + \beta_2 FL_{jt}$$

In these equations the index j refers to individual OECD countries (j=1,...,15 or 19) and the index t to time (t=1990,...1999). As we have mentioned before, b_t and FL_t refer to the beginning of period t. Following the majority of studies on the effects of financial deregulation and liberalisation, one may expect that $\beta_2<0$. The sign to be expected for β_1 is unclear. If the liquidity constraints effect of government debt dominates, β_1 should be positive. If the effect on myopia dominates, it should be negative. Substituting (6) into (5), it follows that:

$$(7)\ \Delta c_{jt} = \alpha_j + \beta_0 \Delta y_{jt} + \beta_1 b_{jt} \Delta y_{jt} + \beta_2 FL_{jt} \Delta y_{jt} + \varepsilon_{jt}$$

3.2. Preliminary econometric considerations

Data and data sources.

Table 1 below describes our data and data sources. We use annual data for the OECD countries in the 1990s. All data are standard, except our indicator for financial liberalisation. Although we will consider alternative indicators in section 3.4., our main results include the number of credit cards issued by Visa (see also Callen and Thimann, 1997). This approach deviates from existing work, where typically the stock of outstanding consumer credit is used (e.g. Girardin et al., 2000; Bachetta and Gerlach, 1997). For our purpose this variable is inadequate because it may be highly endogenous to the evolution of government debt, which is another variable in (6). Indeed, as we have argued before, one of the reasons for the government debt ratio to affect the excess sensitivity of private consumption to disposable income may be that it makes banks less willing to lend. We prefer the number of credit cards outstanding because it can reasonably be expected to be more of an exogenous nature than other variables.

To estimate equation (7) we employ two alternative series for the government debt ratio. Interestingly, if a higher government debt mainly operates by affecting the banks' willingness to lend (liquidity constraints effect), the net debt ratio might be the more relevant variable. To calculate households' future tax liabilities one can expect rational banks to take into account the government's financial assets. On the other hand, if a higher government debt mainly operates by alerting unconscious, myopic consumers, the gross government debt ratio might be more relevant. If government debt is discussed in politics or the media, the numbers typically refer to gross government debt.

Table 1. Data and data sources.

C_{jt}	Private consumption in real per capita terms. Available from OECD Statistical Compendium on CD-rom (2001-II). Available for all countries.
Y_{jt}	Household disposable income in real per capita terms. Deflated by index for private consumption. Available from OECD. Available for all countries except Greece. For Greece we proxy disposable income through GDP minus net taxes. The latter are calculated as the sum of government consumption and government savings, both available from OECD.
bg_{jt}	Ratio of gross government debt to GDP. Available in 19 OECD countries: Australia, Greece, Ireland, Portugal and the 15 countries that occur in figure 1. For Finland observations are only available starting in 1989.
bn_{jt}	Ratio of net government debt to GDP. Available in each of the 15 countries that occur in figure 1. Not available for Australia, Greece, Ireland and Portugal.
FL_{jt}	Per capita number of credit cards issued by Visa International (Visa, inc.). Available in all 19 countries since 1989 (or earlier for some countries).

Note: Since the theoretical variables bg_{jt}, bn_{jt} and FL_{jt} refer to the beginning of year t, we will in our empirical work use data for the previous year.

Estimation method.

Empirically, a number of econometric issues and complications have to be dealt with. First, Δy_{jt} being correlated to shocks in consumption (ε_{jt}), an instrumental variables approach is needed. Second, as shown by Campbell and Mankiw (1990), variation of λ over time and across countries implies heteroskedasticity in the error term ε_{jt}. Another element of cross-country variation concerns the unobserved country-specific effects α_j. Moreover, given that consumption growth is a major component of output and income growth in macroeconomic data, correlation between α_j and Δy_{jt} is obvious. An appropriate way to deal with these problems of endogeneity and heterogeneity is the use of GMM after first-differencing equation (7)[4]. A third complication is related to time aggregation in available consumption and income data. Theoretically, this induces an MA(1) component in ε_{jt} (Working, 1960). So does the inclusion of expenditures on durables in our measure of consumption (Mankiw, 1982)[5]. Consistent estimation would then require an instrument set with at least a two period gap between the regressors and the instruments. Our first-differenced GMM approach reinforces this problem. Since this approach comes down to estimating an equation for the second difference of consumption, it implies an MA(2) process in the error term. Reliable instruments should then be lagged three times. Weak forecasting power is to be expected. The use of weak instruments may result in biased coefficients in small samples (see Loayza et. al., 2000). Empirically, the problem need not be that big though. For example, also estimating a consumption function, Lopez et al. (2000) cannot reject the hypothesis that ε_{jt} does *not* contain an MA(1) component in a panel of 19 OECD countries. Actually, this is also what we shall find (see below). Twice lagged instruments are in that case reliable. Further details about these instruments are discussed in the next section.

3.3. Empirical results

Table 2 presents our main results. As a measure for government debt we use the net debt ratio. Data are available for 15 countries in 1990-99 (see table 1). In table 4 we include the gross government debt ratio. Although data are then available for more countries, for reasons

[4] See e.g. Bond (2002) for a general discussion. For an excellent description of this method applied to macroeconomic consumption or savings data, see Loayza et al. (2000) and Lopez et al. (2000).

[5] Furthermore, an MA(1) component in ε_{jt} may show up if consumption levels contain a transitory component.

to be discussed below we consider the results in table 4 to be somewhat less reliable. As shown by Hansen (1982), the optimal GMM estimator is obtained in two steps. In our discussion, we focus on the second-step results[6]. On the whole, the specification tests in table 2 (Sargan test for overidentifying restrictions and tests for first order and second order serial correlation) do not show evidence against our estimates. The absence of significant second order serial correlation justifies our use of twice lagged 'internal' instruments (shown at the bottom of table 2)[7]. The Sargan test does not reject their joint validity. Regarding the point estimates, regression (R1) imposes the restriction that the degree of excess sensitivity is constant over time and across countries. The result confirms the existing literature. Rejecting Hall (1978), it is consistent with the hypothesis that a significant fraction of consumers is either liquidity constrained or myopic. Regression (R2) demonstrates the time-varying nature of this fraction. A new result is that this fraction is significantly affected by the level of the (net) government debt ratio. The positive sign of β_1 suggests that a higher government debt ratio raises the excess sensitivity of consumption because it reinforces the incidence of liquidity constraints, a result that is fully consistent with the relationship depicted in figure 1. Any effect of a higher government debt ratio on the fraction of myopic consumers, if it exists, is clearly dominated by the opposite effect on the fraction of credit constrained consumers. Regressions (R3) and (R4) introduce financial liberalisation as an explanatory variable. As we have mentioned before, we follow Callen and Thimann (1997) using data on the per capita number of credit cards issued by Visa International. In contrast to many other studies (see section 1), we find no significant negative effect on excess sensitivity from (our proxy for)

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[6] Bond (2002) argues that the asymptotic standard errors of the two-step GMM estimates may be a poor guide for hypothesis testing in certain cases. As noted by Bond and Windmeijer (2002) this problem is especially relevant when the number of instruments grows rapidly with the time dimension, which is not the case here since we choose a fixed number of instruments per time period (for reasons explained in Loayza et. al. 2000). We nevertheless report both the first and second step results.

[7] Using instruments that are lagged three times also yields insignificant results for second order serial correlation in the error term. Although with these instruments - most likely due to their weak forecasting power - the precision of our estimates is affected, our main conclusions about the signs and significance of the parameters in table 2 are unaffected (results available upon request).

financial deregulation. The net government debt ratio however maintains a significant positive coefficient in regression (R4).

Table 2. Estimation results for equation (7) using the net government debt ratio, 1990-99 [a]

	One-step estimates with heteroskedasticity consistent standard errors			
Estimated parameter	(R1)	(R2)	(R3)	(R4)
β_0	0.35 (3.24)	0.24 (2.35)	0.34 (2.13)	0.29 (1.61)
β_1	-	0.48 (2.80)	-	0.41 (2.55)
β_2	-	-	0.12 (0.18)	-0.26 (0.35)
	Two-step GMM estimates			
Estimated parameter	(R1)	(R2)	(R3)	(R4)
β_0	0.35 (10.6)	0.23 (5.37)	0.40 (3.10)	0.30 (2.53)
β_1	-	0.63 (6.27)	-	0.56 (5.14)
β_2	-	-	-0.20 (0.31)	-0.37 (0.56)
N. Obs.	150	150	150	150
Sargan (p-value) [(b)]	0.60	0.98	0.95	0.99
Test for first order serial correlation (p-value) [(c)]	0.016	0.048	0.021	0.054
Test second order serial correlation (p-value) [(c)]	0.191	0.295	0.249	0.397
Instrument set	Δc_{jt-2}, Δy_{jt-2}	Δc_{jt-2}, Δy_{jt-2}, $bn_{jt-2}\Delta y_{jt-2}$	Δc_{jt-2}, Δy_{jt-2}, $FL_{jt-2}\Delta y_{jt-2}$	Δc_{jt-2}, Δy_{jt-2}, $bn_{jt-2}\Delta y_{jt-2}$, $FL_{jt-2}\Delta y_{jt-2}$

Notes: [a] Absolute t-statistics in parentheses; [b] Sargan is Sargan test of overidentifying restrictions. The null hypothesis is that the overidentifying restrictions are correct; [c] The null hypothesis is that there is no first (second) order serial correlation in the error term.

There are various possible explanations for the insignificance of the degree of financial liberalisation. A first one may be that this paper only investigates the 1990s, whereas earlier studies also included earlier decades. This explanation may make sense. It suggests that after a decade of liberalisation in the 1980s, the effect on bank lending to consumers of further liberalisation in the 1990s weakened. Another reason may be our panel approach. For example, McKiernan (1996) and Girardin et al. (2000) only studied one particular country. Clearly, if this were the reason, our results have the advantage of being more general. A third explanation might be that our proxy for financial liberalisation is inadequate. In that case, our results would be biased against the hypothesis that financial deregulation matters for private consumption and its responsiveness to current income. To assess the relevance of this potential problem, we employ several alternative proxies in section 3.4.

Using the results of regression (R4), table 3 shows the estimated excess sensitivity parameter λ_{jt} for each country in the beginning and at the end of the 1990s. Countries are ranked according to their net government debt ratio in 1989. Unsurprisingly, given the highly significant and positive β_1 in table 2, we observe typically higher excess sensitivity in the highest debt countries (Belgium, Italy) and typically lower excess sensitivity in the lowest debt countries (Finland, Norway,...). Another interesting observation is that, on average, the estimated λ_{jt} hardly changed between the beginning and the end of the 1990s. It was close to 0.40, both in 1990 and in 1999. This is surprising, observing that the (unweighted) average net government debt ratio in the countries included in table 3 increased from about 30% in 1989 to more than 40% in 1998. The obvious explanation for this paradox concerns further financial liberalisation. In most countries the per capita number of Visa cards rose strongly in the 1990s. However, given the statistically insignificant result for Visa in table 2, we should be cautious in drawing this conclusion.

Table 3. Estimated excess sensitivity (λ_{jt}), using the results of regression (R4) in table 2.

	1990	1999		1990	1999
Net debt > 50% of GDP in 1989			Net debt < 25% of GDP in 1989		
Belgium	0.92	0.85	Finland	0.06	0.06
Italy	0.80	0.84	France	0.33	0.45
Canada	0.44	0.48	Germany	0.40	0.53
			Japan	0.31	0.28
Net debt between 25% and 50% of GDP in 1989			Norway	0.02	-0.12
			Sweden	0.27	0.26
Austria	0.50	0.53	UK	0.26	0.24
Denmark	0.47	0.41			
Netherlands	0.48	0.57			
Spain	0.40	0.39	All country average (unweighted)	0.403	0.395
US	0.39	0.16			

Table 4 presents estimation results for 19 OECD countries using the gross government debt ratio.

Table 4. Estimation results for equation (7) using the gross government debt ratio, 1990-99 [a]

	One-step estimates with heteroskedasticity consistent standard errors			
Estimated parameter	(R1)	(R2)	(R3)	(R4)
β_0	0.30 (3.30)	0.06 (0.32)	0.24 (1.99)	0.15 (0.58)
β_1	-	0.45 (1.71)	-	0.38 (1.33)
β_2	-	-	0.40 (0.76)	-0.17 (0.28)
	Two-step GMM estimates			
Estimated parameter	(R1)	(R2)	(R3)	(R4)
β_0	0.28 (10.85)	0.09 (1.10)	0.30 (4.22)	0.23 (2.12)
β_1	-	0.42 (4.52)	-	0.44 (2.76)
β_2	-	-	0.02 (0.07)	-0.73 (1.69)
N. Obs.	189	189	189	189
Sargan (p-value) [(b)]	0.39	0.80	0.85	0.99
Test for first order serial correlation (p-value) [(c)]	0.006	0.008	0.011	0.008
Test second order serial correlation (p-value) [(c)]	0.111	0.065	0.115	0.080
Instrument set	Δc_{jt-2}, Δy_{jt-2}	Δc_{jt-2}, Δy_{jt-2}, $bg_{jt-2}\Delta y_{jt-2}$	Δc_{jt-2}, Δy_{jt-2}, $FL_{jt-2}\Delta y_{jt-2}$	Δc_{jt-2}, Δy_{jt-2}, $bg_{jt-2}\Delta y_{jt-2}$, $FL_{jt-2}\Delta y_{jt-2}$

Notes: [a] Absolute t-statistics in parentheses; [b] Sargan is Sargan test of overidentifying restrictions. The null hypothesis is that the overidentifying restrictions are correct; [c] The null hypothesis is that there is no first (second) order serial correlation in the error term.

For two reasons we believe that these results may be somewhat less reliable, despite the larger sample size. First, our findings up to now strongly suggest dominance of the liquidity constraints effect of rising government debt. If this conclusion were confirmed (and it will), the gross government debt ratio may be not the most appropriate variable. As we have mentioned before, in their lending decisions one may expect rational banks to take into account net rather than gross government debt. The test results for second order serial correlation in the error term are another reason for caution. Although the null hypothesis of no second order correlation can never be rejected at the 5% level, it sometimes can at the 10% level. To be on the safe side, it might then be preferable to adopt a three times lagged instrument set. However, the forecasting power of these instruments being much lower, we would have to pay the price of biased and imprecise estimates. We have therefore chosen to stick to a twice lagged instrument set. Discussion of the (two-step) point estimation results in table 4 can be brief. They fully confirm those of table 2. We always obtain a positive and statistically significant β_1 and an insignificant (or only marginally significant) β_2. In regression (R3) β_2 even has the wrong sign.

3.4. Robustness tests

In this section we organise three robustness checks on our results. The first one introduces different proxies for financial liberalisation. The second one allows for unobserved country-specific fixed determinants of the degree of excess sensitivity (equation 6). The third robustness check allows for a richer set of effects from the government debt ratio than specified in equations (5) and (6).

Table 5 presents the results from re-estimating equation (7) with the net government debt ratio, but using three alternative indicators for financial liberalisation. These are (i) the ratio of nominal M2 to GDP, (ii) the ratio of nominal M1 to M2 and (iii) the trend in the spread between the banks' lending and deposit rates. We also mention the included proxies at the bottom of table 5[8]. Appendix A describes the correlation between them. Interestingly, the majority of pairwise correlation coefficients is smaller than 0.33 in absolute value. This obviously strengthens the case for robustness checks. Also, it raises the power of our results if they survive these checks.

[8] The trend in (iii) has been obtained from a Hodrick-Prescott filter. Data for M1 and M2 in most countries have been taken from OECD, Statistical Compendium on CD-rom, 2001-II. For the UK and Sweden data for M1 are not available. As a proxy we use the money base, taken from IMF, International Financial Statistics. Data for the banks' lending and deposit rates have also been taken from IMF, International Financial Statistics. For Austria these data are not available, for the UK one observation is missing, which explains the lower number of observations in regression (R3) in table 5.

Since we expect M2/GDP to be positively related to the degree of financial liberalisation, this variable should get a negative sign in equation (7). For M1/M2 and the interest rate spread the opposite applies. The results in table 5 are rather disappointing. Again concentrating on the second step estimates, β_2 obtains the wrong sign in (R1) and (R2). In (R1) the positive β_2 is even statistically signficant. In (R3) β_2 obtains the expected sign, but it is highly insignificant. The conclusion seems to be unavoidable that there is no evidence that financial liberalisation contributed to a lower degree of excess sensitivity of private consumption in the 1990s. By contrast, our findings for the government debt ratio survive this first robustness check, especially in regressions (R2) and (R3). In (R1) β_1 is statistically significant only at the 10% level (second step estimate).

Table 5. Estimation results for equation (7) using the net government debt ratio and alternative proxies for financial liberalisation, 1990-99 [a]

	One-step estimates with heteroskedasticity consistent standard errors		
Estimated parameter	(R1)	(R2)	(R3)
β_0	-0.21 (0.56)	0.36 (1.70)	0.20 (1.32)
β_1	0.29 (1.39)	0.44 (2.18)	0.43 (1.91)
β_2	0.94 (1.53)	-0.31 (0.94)	-0.0001 (0.004)
	Two-step GMM estimates		
Estimated parameter	(R1)	(R2)	(R3)
β_0	-0.36 (1.05)	0.57 (1.87)	0.15 (1.25)
β_1	0.35 (1.89)	0.50 (2.87)	0.56 (4.51)
β_2	1.23 (2.12)	-0.67 (1.23)	0.009 (0.34)
N. Obs.	150	150	139
Sargan (p-value) [(b)]	0.99	0.99	0.99
Test for first order serial correlation (p-value) [(c)]	0.009	0.035	0.088
Test second order serial correlation (p-value) [(c)]	0.130	0.163	0.212
Instrument set	Δc_{jt-2}, Δy_{jt-2}, $bn_{jt-2}\Delta y_{jt-2}$, $FL_{jt-2}\Delta y_{jt-2}$	Δc_{jt-2}, Δy_{jt-2}, $bn_{jt-2}\Delta y_{jt-2}$, $FL_{jt-2}\Delta y_{jt-2}$	Δc_{jt-2}, Δy_{jt-2}, $bn_{jt-2}\Delta y_{jt-2}$, $FL_{jt-2}\Delta y_{jt-2}$
Included proxy for *FL*	M2/GDP	M1/M2	Spread between lending and deposit rates (trend)

Notes: [a] Absolute t-statistics in parentheses; [b] Sargan is Sargan test of overidentifying restrictions. The null hypothesis is that the overidentifying restrictions are correct; [c] The null hypothesis is that there is no first (second) order serial correlation in the error term.

The regressions in table 6 allow for the possibility of unobserved country-specific fixed determinants of λ. More precisely, we include country dummies in equation (6) which should capture the influence of unknown or hard to measure institutional or structural differences across countries that affect the degree of excess sensitivity. Haliassos and Christou (2000) for example point to differences related to the structure of the banking industry that may affect lending policy and the incidence of liquidity constraints. To the extent that consumer wealth is related to age (e.g. Kennickell and Starr-McCluer, 1997), the age structure of the population may also matter for the incidence of liquidity constraints. Another issue concerns different consumer preferences for Visa versus other credit card institutions. If there are structural differences in these preferences across countries, the evolution of the number of Visa cards may be a good proxy for financial liberalisation *in each country over time*, but their levels in a particular year may not be comparable *across countries*.

Ideally, country dummies are included for all countries, except one. Doing this, however, one runs into problems. Including dummies for all countries in equation (6) implies an enormous increase in the number of slope coefficients to be estimated on Δy in (7). The problem is that as the number of explanatory variables in that equation rises, so do the standard errors on all estimated coefficients, including those on the government debt ratio[9]. Gujarati (2003) also points to possible problems of imprecise estimation (multicollinearity) in panels when too many dummy variables are included. Underlying the results in table 6 (R1, R2, R3) is an alternative approach. We have added dummies for each country separately. Insignificant dummies were dropped, whereas dummies that showed up significant at 10% or better were kept in the regression. Any time a dummy "survived", we re-tested the significance of the

[9] Another, even more serious problem occurs if one wants to include twice lagged values of the additional explanatory variables among the set of instruments. Given the relatively large time dimension of our sample we would end up in a situation with more moment conditions (instruments) than observations, which is clearly not feasible.

others. In the end we have found significantly different values for β_0 only in Denmark, Germany and the UK.

Table 6. Estimation results for equation (7) allowing for differences in β_0 and β_1 across countries, 1990-99 [a]

	One-step estimates with heteroskedasticity consistent standard errors			
Estimated parameter	(R1)	(R2)	(R3)	(R4)
β_0	0.31 (2.22)	0.02 (0.10)	-0.05 (0.23)	0.30 (2.03)
β_1	0.43 (3.53)	0.25 (1.37)	0.61 (1.99)	0.43 (3.30)
β_2	0.52 (0.84)	0.60 (1.32)	0.29 (0.52)	0.13 (0.22)
Countries with signif. different β_0 [(b)]	Denmark (-), Germany (+), UK (-)	Denmark (-), Germany (+), UK (-)	Germany (+), UK (-)	
Countries with signif. different β_1				Denmark (-)
	Two-step GMM estimates			
Estimated parameter	(R1)	(R2)	(R3)	(R4)
β_0	0.43 (1.72)	0.63 (1.18)	0.03 (0.14)	0.42 (2.28)
β_1	0.67 (3.22)	0.72 (2.24)	0.69 (3.72)	0.65 (4.98)
β_2	-0.22 (0.14)	-0.33 (0.36)	-0.30 (0.60)	-0.63 (-0.59)
Countries with signif. different β_0 [(b)]	Denmark (-), Germany (+), UK (-)	Denmark (-), Germany (+), UK (-)	Germany (+), UK (-)	
Countries with signif. different β_1 [(b)]				Denmark (-)
N. Obs.	150	150	189	150
Sargan (p-value) [(c)]	0.99	0.99	0.99	0.99
Test for first order serial correlation (p-value) [(d)]	0.013	0.002	0.006	0.047
Test second order serial correlation (p-value) [(d)]	0.204	0.125	0.162	0.445
Instrument set	Δc_{jt-2}, Δy_{jt-2}, $bn_{jt-2}\Delta y_{jt-2}$, $FL_{jt-2}\Delta y_{jt-2}$, $\Delta y_{jt-2}dum_j$ *(for each j where* β_0 *is different)*	Δc_{jt-2}, Δy_{jt-2}, $bn_{jt-2}\Delta y_{jt-2}$, $FL_{jt-2}\Delta y_{jt-2}$, $\Delta y_{jt-2}dum_j$ *(for each j where* β_0 *is different)*	Δc_{jt-2}, Δy_{jt-2}, $bg_{jt-2}\Delta y_{jt-2}$, $FL_{jt-2}\Delta y_{jt-2}$, $\Delta y_{jt-2}dum_j$ *(for each j where* β_0 *is different)*	Δc_{jt-2}, Δy_{jt-2}, $bn_{jt-2}\Delta y_{jt-2}$, $FL_{jt-2}\Delta y_{jt-2}$, $bn_{jt-2}\Delta y_{jt-2}dum_j$ *(for each j where* β_1 *is different)*
Included variables for *b* and *FL*	net debt ratio, visa	net debt ratio, M2/GDP	gross debt ratio, visa	net debt ratio, visa

Notes: [a] Absolute t-statistics in parentheses; [b] The positive or negative sign behind the name of each country indicates whether the included dummy for this country is negative or positive, that is whether β_0 or β_1 for this country is significantly higher (+) or lower (-) than the reported β_0 or β_1. Country dummies are included if they are significant at 10% or better in the two-step GMM regression; [c] Sargan is Sargan test of overidentifying restrictions. The null hypothesis is that the overidentifying restrictions are correct; [d] The null hypothesis is that there is no first (second) order serial correlation in the error term.

For all other countries we have imposed the restriction that β_0 is identical. As to the estimation results, the first and the second regression in table 6 include the net government debt ratio as an explanatory variable, the third one includes the gross government debt ratio. Further, (R1) and (R3) include the per capita number of Visa cards as a proxy for financial liberalisation, (R2) includes the ratio of M2 to GDP. We pay special attention to the latter variable because of the results in table 5. In that table, the ratio of M2 to GDP obtained a significantly positive (wrong) sign. Moreover, including M2/GDP somewhat affected the statistical significance of the government debt ratio. The results from this second robustness check are reassuring. Again considering the second step estimates, controlling for unobserved country-specific fixed determinants of the degree of excess sensitivity does not at all affect our conclusions. The effect from the government debt ratio (β_1) remains robustly positive and significant. Our proxies for financial liberalisation obtain the correct (negative) sign in each regression, but they are always statistically insignificant. The latter result is especially interesting for the ratio of M2 to GDP. Controlling for unobserved country-specific effects, both the sign and the degree of statistical significance of this variable change.

Given that the value of β_1 may also be different across countries (e.g. because of differences in financial institutions across countries or differences in poverty levels) in (R4) in table 6 we report the results of following the same strategy for β_1. In the end we have found significantly different values for β_1 only in Denmark. For all other countries we have imposed the restriction that β_1 is identical.

The results in table 7 test for additional effects from the government debt ratio in equations (5) or (6). Regressions (R1) and (R2) follow from including the net government debt ratio as an additional variable in equation (5). That is

$$(5') \quad \Delta c_{jt} = \alpha_j + \lambda_{jt}\Delta y_{jt} + \rho b_{jt} + \varepsilon_{jt}$$

The estimated equation (7) then is

$$(7') \quad \Delta c_{jt} = \alpha_j + \beta_0 \Delta y_{jt} + \beta_1 b_{jt}\Delta y_{jt} + \beta_2 FL_{jt}\Delta y_{jt} + \rho b_{jt} + \varepsilon_{jt}$$

A potential justification for this extension of equation (5) relates to the idea of precautionary savings. More precisely, an increase in the government debt ratio may raise uncertainty among consumers. Fiscal policy may become less predictable at high debt rates and consumers' assessment of the variance in their future consumption may rise. Facing more uncertainty, they may prefer to decrease current consumption and increase savings. A steeper consumption path over time will result. Algebraically, this precautionary motive suggests ρ to be positive. The results in (R1) and (R2) fully confirm this expectation: ρ is estimated to be about 0.08 and statistically very significant. Quite important also, including the government debt ratio as a separate variable in (7') does not affect our previous results. We still find that a higher government debt ratio raises the degree of excess sensitivity of private consumption to current income (β_1).

Regressions (R3) and (R4) model the relationship between the degree of excess sensitivity and the government debt ratio in equation (5) as a linear spline. This is a piecewise linear relationship between λ and b with the line segments joining one another at one or more breakpoints. We allow for one breakpoint at a debt ratio equal to 40%. The intercept and the slope (β_1) describing the effect of government debt on excess sensitivity can differ for both segments. Speculating that banks are more likely to restrict private credit when the government debt ratio is considered to be a problem, our intuition is that β_1 should be higher for high debt ratios. Regressions (R3) and (R4) confirm this intuition. For debt ratios below 40%, we do obtain a positive effect on excess sensitivity (β_1 is about 0.3), but is not statistically significant. For debt ratios higher than 40% this positive effect is not only stronger, it is also significant at about 5% or better. Highly similar results are obtained if we choose 30% or 50% as our "critical value".

Table 7. Estimation results allowing for additional effects from the (net) government debt ratio, 1990-99 [a]

	One-step estimates with heteroskedasticity consistent standard errors			
Estimated parameter	(R1)	(R2)	(R3)	(R4)
β_0	0.31 (3.16)	0.43 (3.12)	0.20 (2.20)	0.25 (1.62)
β_1	0.51 (4.34)	0.43 (5.48)	-	-
β_2	-	-0.28 (0.56)	-	-0.31 (0.53)
ρ	0.09 (2.85)	0.10 (3.54)	-	-
β_1 (for $bn_{jt} \leq 0.4$)	-	-	0.27 (0.57)	0.23 (0.51)
β_1 (for $bn_{jt} > 0.4$)	-	-	0.47 (2.15)	0.45 (1.98)
	Two-step GMM estimates			
Estimated parameter	(R1)	(R2)	(R3)	(R4)
β_0	0.33 (2.83)	0.49 (3.98)	0.18 (3.23)	0.27 (1.94)
β_1	0.57 (4.01)	0.44 (2.92)	-	-
β_2	-	-0.24 (0.66)	-	-0.53 (0.83)
ρ	0.08 (3.41)	0.08 (3.80)	-	-
β_1 (for $bn_{jt} \leq 0.4$)	-	-	0.34 (1.26)	0.30 (1.10)
β_1 (for $bn_{jt} > 0.4$)	-	-	0.79 (2.15)	0.76 (1.94)
N. Obs.	150	150	150	150
Sargan (p-value) [(b)]	0.99	0.99	0.99	0.99
Test for first order serial correlation (p-value) [(c)]	0.015	0.005	0.069	0.08
Test second order serial correlation (p-value) [(c)]	0.265	0.191	0.405	0.624
Instrument set [(d)]	Δc_{jt-2}, Δy_{jt-2}, $bn_{jt-2}\Delta y_{jt-2}$, bn_{jt-2}	Δc_{jt-2}, Δy_{jt-2}, $bn_{jt-2}\Delta y_{jt-2}$, bn_{jt-2}, $FL_{jt-2}\Delta y_{jt-2}$	Δc_{jt-2}, Δy_{jt-2}, $(bn*dum+0.4(1-dum))_{jt-2}\Delta y_{jt-2}$ $((bn-0.4)*(1-dum))_{jt-2}\Delta y_{jt-2}$	Δc_{jt-2}, Δy_{jt-2}, $FL_{jt-2}\Delta y_{jt-2}$, $(bn*dum+0.4(1-dum))_{jt-2}\Delta y_{jt-2}$ $((bn-0.4)*(1-dum))_{jt-2}\Delta y_{jt-2}$
Included variable for *FL*	-	Visa	-	visa

Notes: [a] Absolute t-statistics in parentheses; [b] Sargan is Sargan test of overidentifying restrictions. The null hypothesis is that the overidentifying restrictions are correct; [c] The null hypothesis is that there is no first (second) order serial correlation in the error term. [d] dum is a dummy variable that equals 1 if bn< 0.4 and 0 otherwise.

4. Conclusions.

Empirical studies have typically rejected Hall's (1978) conclusion that the change in private consumption cannot be forecasted. They have shown that private consumption is much more sensitive to changes in current disposable income than predicted by Hall. Standard explanations for this "excess sensitivity" of private consumption refer to liquidity constraints and/or myopia. Building on the idea of liquidity constraints, several authors have more recently endogenised the degree of "excess sensitivity". Some find that countries with better developed credit markets have a smaller fraction of liquidity constrained consumers and therefore a smaller degree of excess sensitivity. In line with this result, a number of studies find that when countries go through a process of financial liberalisation, the degree of excess sensitivity falls over time. Other studies however cannot confirm that conclusion.

In this paper we demonstrate the crucial role of the government debt ratio for the degree of excess sensitivity of private consumption to current income changes. Theoretically, this role can be rationalised from different angles. On the one hand, a high or (rapidly) increasing government debt ratio may alert unconscious citizens and, as a consequence, reduce the fraction of myopic consumers. Excess sensitivity of private consumption to current income should then fall. On the other hand, high or rising government debt ratios imply an increase in households' future liabilities. Banks may then reduce the amounts they lend, thereby raising the incidence of liquidity constraints and excess sensitivity of private consumption. Empirically, we test the relevance of these hypotheses in a panel data study of private consumption in the OECD countries in the 1990s. To assess the influence of the government debt, we include both the net and the gross debt ratio. Furthermore, to estimate the effects of financial liberalisation we include several proxies. Our results strongly support the idea that a higher government debt ratio implies tighter credit conditions and an increase of excess sensitivity. This unfavourable effect from government debt on credit conditions seems to be strong especially when government debt is already at a high level. Any effect of a higher government debt ratio on the fraction of myopic consumers, if it exists, is clearly dominated by the opposite effect on the fraction of credit constrained consumers. As to the effects of financial liberalisation, we find no convincing evidence that it reduced excess sensitivity, at least not in the 1990s.

What are the implications of our findings? First of all, our results suggest that fiscal policy may be more effective (Keynesian) at high debt rates. As is well known from macroeconomics textbooks, the responsiveness of private consumption to current income is a crucial determinant of the Keynesian multiplier. We show that at high debt rates the responsiveness of private consumption (and by consequence the multiplier) may be significantly higher. In this respect, the policy implication of our findings may differ from those of e.g. Nicoletti (1988), Sutherland (1997) or Zaghini (1999), who all suggest that at high debt rates fiscal policy may become less effective. Note however that to the extent that lenders tighten credit conditions when the government debt rises, fiscal impulses may also have negative effects. Tax reductions financed by bonds are not only good news for liquidity constrained consumers (who can spend more), they may also be bad news for initially unconstrained consumers who cannot borrow anymore (or can only borrow less) because banks tighten credit conditions. These consumers may then have to postpone, say, the purchase of a home or a durable, which undermines the effectiveness of fiscal policy.

Obviously, although the potentially increased effectiveness of fiscal policy at high debt rates may be good news from the perspective of stabilisation policy, this does not mean that it is welfare improving. The reason for higher effectiveness is precisely that more consumers are liquidity constrained, and therefore prevented from realising their optimal consumption path.

As a second implication, our findings reveal a weakness in recent studies investigating the effects of fiscal policy on private consumption and savings (e.g. Evans and Karras, 1998; Perotti, 1999; Lopez et al., 2000). All these studies assume the existence of two or three groups of consumers, the fractions of which are considered constant (e.g. permanent income consumers, liquidity constrained consumers, etc.). Our results challenge this assumption. The stance of fiscal policy itself and the government debt ratio may change these fractions.

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Appendix A: Correlation over all countries and years between proxies for financial liberalisation .

Table A1. Pairwise correlation matrix.

	Visa	M2/GDP	M1/M2	Spread (trend)
Visa	1.000			
M2/GDP	0.066	1.000		
M1/M2	-0.307	-0.524	1.000	
Spread (trend)	-0.719	-0.291	0.243	1.000

Note : All correlations are based on data for 1990-99 in 15 OECD countries, except correlations with the spread between lending and deposit rates. For this variable data are not available for Austria.