CENTRE
D'ÉTUDES PROSPECTIVES
ET D'INFORMATIONS
INTERNATIONALES

No 2007 – 10
May

Fiscal Policy in Real Time

Jacopo Cimadomo

Fiscal Policy in Real Time

Jacopo Cimadomo

No 2007 – 10
May

Contents

FISCAL POLICY IN REAL TIME

NON-TECHNICAL SUMMARY

Most of the empirical literature on fiscal policy has found that, over the post-World War II period, governments in developing and industrialized countries have reacted "pro-cyclically" to fluctuations in the economic activity (see e.g. Lane (2003) and Kaminsky, Reinhart and Vegh (2004)). Otherwise stated, budgetary decisions such as tax increases and cuts in public spending implemented in "bad times" would have tended to aggravate the length and the severity of economic recessions. On the other side, expansive policies put in place during "good times" would have led economic booms to be more prolonged and vigorous.

This empirical evidence has been mainly drawn from the estimation of fiscal policy reaction functions, relating a policy indicator to the output gap and other explanatory variables, based on the use of revised data, i.e. data available in an "updated" form to the econometrician at the time the study is carried out. Since many economic variable are seriously contaminated by revision errors, however, revised data may be substantially different from the ones available in "real-time" to policymakers at the time of budgeting. In other words, as shown by Orphanides (2001) in the framework of monetary policy analysis, unrealistic assumptions about the timeliness of data availability may induce misleading assessments on the historical policy stance. Nevertheless, although informational problems clearly matter also for the evaluation of the fiscal policy stance, little has been done in this field.

In the present study we show that, when the object of interest is *intentional* stance of fiscal policy, real-time information on *all* the variables included in a fiscal policy rule should be employed. In particular, it is highlighted that the use of real-time observations on the fiscal policy "instrument" itself, typically the structural primary balance, may be of key importance. In fact, and in contrast with central bankers who can control their operating instrument, the short-term interest rate, with great precision, the actual realization of planned fiscal measures may depend on several factors outside the direct control of fiscal authorities. Hence, there might be sizeable differences between discretionary fiscal measures as planned in the past and what it is observed *ex-post*, for the same years.

Based on a dataset of revised and real-time observations drawn from the December Issues of the OECD Economic Outlook for 19 industrialized countries, from 1994 to 2006, it is shown that the stance of fiscal policy seems to be pro-cyclical, if evaluated *ex-post*. When real-time data are used in the estimation of fiscal policy rules, however, the *ex-ante* stance appears to be counter-cyclical, especially during buoyant economic times. The analytical form of the bias incurred in evaluating the *intentional* stance of the policy using revised data is formally derived. It is demonstrated that the size and the sign of that bias can be accurately predicted, based on empirical second-order moments of revisions errors in the variables of interest.

Finally, the possible presence of non-linearities in the way the discretionary component of fiscal policy reacts to the economic cycle and debt accumulation is tested. It emerges that the *intentional* behavior of fiscal policy is characterized by two regimes, and that the switch between them, from a neutral or slightly pro-cyclical stance to a counter-cyclical one, is likely to occur when output is close to its potential level. However, the hypothesis of threshold effects is always rejected when the analysis is based on revised data.

ABSTRACT

In this paper we argue that any assessment on the *intentional* stance of fiscal policy should be based upon *all* the information available to policymakers at the time of fiscal planning. In particular, real-time data on the discretionary fiscal policy "instrument", the structural primary balance, should be used in the estimation of fiscal policy reaction functions. In fact, the *ex-post* realization of discretionary fiscal measures may end up to be drastically different from what intentionally planned by fiscal authorities in the budget law. If this is the case, and if revision errors in the policy indicator are correlated with the ones in the regressors, it is shown that commonly used estimators become biased possibly inducing a misleading judgement on the policy stance. We derive the functional form of that bias and, based on empirical second-order moments, we are able to accurately predict the potential impact of using revised data in the evaluation of the *ex-ante* stance of fiscal policy. When fiscal policy rules are estimated on real-time data, our results indicate a counter-cyclical stance in OECD countries, especially during economic expansions. This contrasts with conventional findings based on revised data, which point to fiscal policy acyclicality or pro-cyclicality, and with Forni and Momigliano (2005) who employ real-time data for the output gap and find counter-cyclicality, but just in recessions. Further, we test whether threshold effects might be at play in the reaction of fiscal policy to the economic cycle and to debt accumulation. It emerges that the *intentional* cyclical behavior of fiscal policy is characterized by two regimes, and that the switch between them is likely to occur when output is close to its equilibrium level. On the other hand, the use of revised data does not allow to identify any threshold effect.

JEL Classification: C23, E30, E62, H30
Keywords: Fiscal policy, Cyclical stabilization, Real-time data, Revision errors, Endogenous threshold models.

POLITIQUE BUDGÉTAIRE EN TEMPS REEL

RÉSUMÉ NON TECHNIQUE

Différents travaux récents ont constaté qu'au cours de la période d'après-guerre, les gouvernements des pays en développement comme ceux des pays industrialisés ont réagi de façon "pro-cyclique" aux fluctuations de l'activité économique (voir par exemple Lane (2003) et Kaminsky, Reinhart and Vegh (2004)). Ainsi, des décisions budgétaires telles que des augmentations d'impôts ou des réductions des dépenses publiques mises en place en période de récession ont eu tendance à aggraver la durée et la sévérité des crises. De même, des politiques expansionnistes pratiquées durant des périodes de forte croissance auraient amplifié les reprises.

Ce résultat empirique provient essentiellement de l'estimation de fonctions de réaction des politiques budgétaires, qui relient un indicateur de ces politiques à l'écart de production et à d'autres variables explicatives, toutes basées sur des données révisées. Ces données sont effectivement disponibles pour l'économètre au moment de l'étude. Cependant, beaucoup de variables économiques étant sérieusement affectées par des erreurs de révision, les données révisées peuvent être largement différentes de celles dont les autorités budgétaires disposent en "temps-réel" au moment où elles prennent leurs décisions. En d'autres termes, comme l'a montré Orphanides (2001) dans le cadre de l'analyse de la politique monétaire, des hypothèses peu réalistes sur les données disponibles au moment de la prise de décision peuvent conduire à des évaluations erronées de l'orientation des politiques économiques. Ces problèmes "informationnels" ont été jusqu'ici peu abordés dans l'analyse de la politique budgétaire.

Dans la présente étude, nous montrons que lorsqu'on s'intéresse à l'orientation intentionnelle de la politique budgétaire, c'est l'information en temps-réel sur toutes les variables incluses dans la règle de politique budgétaire qui doit être utilisée. En particulier, l'utilisation des observations en temps-réel concernant l'"instrument" de la politique budgétaire, à savoir le solde structurel primaire, peut être d'une importance cruciale. En fait (et contrairement aux banquiers centraux qui peuvent contrôler avec une grande précision leur instrument opérationnel, le taux d'intérêt à court terme), la mise en œuvre effective des mesures budgétaires approuvées peut dépendre de plusieurs facteurs indépendants de la volonté des autorités budgétaires. Par conséquent, un écart important peut survenir entre les mesures fiscales discrétionnaires réellement décidées et ce qu'observe ex-post, sur les mêmes années, un économètre utilisant des données révisées.

Notre base de données comporte des observations révisées et en temps-réel pour 19 pays industrialisés, provenant des éditions de décembre des Perspectives Économiques de l'OCDE sur les années 1994 à 2006. L'orientation de la politique budgétaire semble être en moyenne pro-cyclique, si évaluée a posteriori. Néanmoins, lorsque les données en temps-réel sont utilisées dans l'estimation, l'orientation intentionnelle de la politique budgétaire apparaît contra-cyclique, particulièrement durant les périodes d'expansion économique. L'évaluation de l'orientation intentionnelle de la politique budgétaire est donc biaisée lorsqu'on se réfère aux données révisées. Nous montrons alors que la taille et le signe de ce biais peuvent être prévus de façon précise à partir des propriétés statistiques des erreurs de révision.

Enfin, la présence de non-linéarités dans la manière dont la composante discrétionnaire de la politique budgétaire réagit au cycle économique et à l'accumulation de la dette publique est examinée. Nos résultats suggèrent que l'orientation intentionnelle de la politique budgétaire est caractérisée par deux régimes, et que le passage d'un régime neutre ou légèrement pro-cyclique à un régime contra-cyclique, peut se produire lorsque l'économie est proche de son niveau potentiel. Ces effets de seuil n'apparaissent pas lorsqu'on utilise les données révisées à la place des données en temps-réel.

RÉSUMÉ COURT

Dans cet article nous montrons que l'évaluation de l'orientation intentionnelle de la politique budgétaire devrait être basée sur toute l'information dont disposent les autorités au moment de la planification budgétaire. En particulier, les données en "temps-réel" sur l'indicateur discrétionnaire de la politique devraient être employées dans l'estimation des fonctions de réaction budgétaires. En effet, la réalisation ex-post des mesures budgétaires discrétionnaires peut se révéler fortement différente de ce qui a été prévu par les gouvernements dans la loi de budget. Si tel est le cas, et si les erreurs de révision de l'indicateur de politique sont corrélées avec celles des variables explicatives, nous démontrons que les estimateurs généralement utilisés sont biaisés. Cela peut induire un jugement erroné sur l'orientation de la politique. Nous dérivons la forme fonctionnelle de ce biais. Ensuite, à partir des moments empiriques de second ordre, nous estimons très précisément l'impact potentiel de l'utilisation des données révisées dans l'évaluation de l'orientation intentionnelle de la politique. Lorsque les règles de politique budgétaire sont estimées sur des données en temps-réel, nos résultats indiquent une orientation contra-cyclique dans les pays de l'OCDE, en particulier durant les phases d'expansion. Ceci contraste avec les résultats conventionnels basés sur les données révisées, qui indiquent un caractére soit neutre soit pro-cyclique de la politique, ainsi qu'avec les résulats de Forni et Momigliano (2005) qui, utilisant des données en temps-réel pour l'écart de production, trouvent une orientation contra-cyclique, mais seulement dans les périodes de récession. Enfin, nous examinons l'existence d'effets de seuil dans la réaction de la politique budgétaire au cycle économique et à l'accumulation de la dette publique. L'orientation cyclique intentionelle de la politique budgétaire apparaît caractérisée par deux régimes. Le passage d'un régime à l'autre se produit lorsque la croissance est proche de son niveau potentiel. Cependant, l'utilisation des données révisées ne permet pas d'identifier ces effets de seuil.

Classification *JEL*: C23, E30, E62, H30
Mots-clé: Politique budgétaire, Stabilisation cyclique, Données en temps-réel, Erreurs de révision, Modèles à seuil endogène .

FISCAL POLICY IN REAL TIME[1]

Jacopo Cimadomo[2]

1 Introduction

The active use of fiscal policy to fine tune the business cycle has not ceased to be a controversial issue among economists. The traditional Keynesian school generally recommends that governments should actively operate to smooth economic fluctuations. In particular, during phases of weak economic growth, they should adopt measures, such as tax cuts or new public investments, to foster a recovery in the economic activity. In contrast, when growth is above potential, they should cut public expenditures or increase taxation. In other words, they should act *counter-cyclically* over the economic cycle.
The Keynesian doctrine has heavily influenced the conduct of economic policies in the post-World War II period. From the 1950s on, and especially during the 1970s and the 1980s, however, Keynesianism was at the center of a very intense debate. In particular, according to economists in the "New Classical" tradition (see Sargent and Wallace (1975), Lucas and Sargent (1978), and more recently Chari and Kehoe (1999)), discretionary fiscal policies may end up to be helpless, or even harmful. In that view, the active use of fiscal policy as a stabilizing tool should be discouraged since: i) recessions might be "self-correcting";[3] ii) there are long and uncertain time lags in the implementation of fiscal measures; iii) institutional constraints may restrict a timely use of fiscal policy; iv) fiscal policy decisions are, often, irreversible.[4]

[1]I especially wish to thank Lucrezia Reichlin for helpful suggestions. I have also benefited from discussions with seminar participants at the Katholieke Universiteit Leuven, CEPII and Paris-Jourdan Sciences Economiques and in particular Manuel Arellano, Agnès Bénassy-Quéré, Martine Carré, Fabrice Collard, Antonello D'Agostino, Domenico Giannone, Luca Sala and Cyrille Schwellnus. This paper was previously circulating under the title "Testing Non-Linearity in Fiscal Policy: New Evidence from Real-Time Data".

[2] CEPII and ECARES-Université Libre de Bruxelles (jacopo.cimadomo@cepii.fr).

[3]New Classical models predict that the market itself takes steps to recover from recessions. In fact, once entrepreneurs realize that a recession is under way, they cut prices to attract new consumers. Workers, in turn, curb their wage demands to reduce unemployment. Thus, the real money supply and aggregate demand automatically rise and, without any government intervention, the output gap shrinks.

[4]The spirit underlying the creation of the European Union fiscal framework, as embedded in the Maastricht Treaty and the Stability and Growth Pact, is to some extent rooted in this debate. The old formulation of the Pact, in fact, suggests that fiscal stabilization should be achieved mainly through the work of automatic stabilizers, once member countries have achieved their medium-term fiscal position

Yet some authors have recently argued that fiscal policy, rather than being counter-cyclical, as Keynesian theories suggest, or acyclical, as advocated by the New Classical macroeconomics, has shown a tendency towards pro-cyclicality.
Among the firsts who explored the issue of pro-cyclicality, Gavin and Perotti (1997) find that fiscal policy in Latin America countries has been characterized as lax during upturns and tight during slowdowns. Kaminsky, Reinhart and Vegh (2004), based on a panel of emerging and industrialized economies, highlight that fiscal policy, but also capital inflows and monetary policy, has been pro-cyclical during the last forty years. Lane (2003) shows that the "pro-cyclicality bias" is more severe in developing countries than in developed ones. As for industrialized countries, the OECD (2003) emphasizes that the stance of fiscal policy tends to be predominantly counter-cyclical in "bad times" and weakly pro-cyclical in "good ones". The European Commission (2004) underlines that discretionary fiscal policies have been mainly pro-cyclical in Euro Area countries throughout the past decades thereby reducing the effectiveness of automatic stabilization: the Stability and Growth Pact would have not helped to eradicate the occurrence of a pro-cyclical bias in these countries. Both the study from the OECD and the one from the European Commission show that sustainability concerns, related to developments in public indebtedness, play a key role in affecting the fiscal stance, fiscal measures becoming tighter when the debt to GDP ratio grows. Summing up, the empirical evidence from this literature seems to be quite consensual as regards developing countries, pointing to a strong pro-cyclicality, whereas results on industrialized economies are more controversial, indicating however some form of pro-cyclicality, in particular during upturns, and especially after 1999.[5]
These studies, though insightful in that they allow to evaluate the *ex-post*, or "realized", stance of fiscal policy, are not suitable to assess the "true", or *intentional*, policy stance since they are based on revised data and not on the information actually available (i.e. available in *real-time*) to policy-makers at the time their decisions have been taken. However, as Orphanides (2001) shows, when unrealistic assumptions on the timeliness of data availability are done, and in particular when it is supposed that the updated, revised information is available *ex-ante* to decision makers, the analysis

of "close-to-balance or in surplus" (see Brunila, Buti and in't Veld (2002)). The new version of the Pact, as from the ECOFIN Council of March 2005, by introducing a more flexible definition of the medium-term objective, implicitly allows discretionary fiscal policies to be used more actively (see Buti and Sapir (2006)).

[5]On the possible factors behind the pro-cyclical bias of the policy, the literature generally distinguishes between emerging and industrial countries. On the former group, for example, Gavin and Perotti (1997) suggest that these countries generally face credit constraints that prevent them from borrowing in bad times. On developed economies, the European Commission (2004) proposes four key factors that can be at the origin of a pro-cyclical fiscal stance: i) uncertainty and measurement errors in potential output and output gap; ii) fiscal rules; iii) political cycles; iv) sustainability concerns.

of policy-makers behavior may be drastically misleading.[6]
Since the seminal work of Orphanides (2001), research employing real-time data has soared in the monetary policy literature (see e.g. Boivin (2005), Croushore and Stark (2001), Giannone, Reichlin and Sala (2005), Ironside and Tetlow (2005)). Although problems related to revisions errors and timeliness of information clearly matter also for the evaluation of the fiscal policy stance, little has been done in the field of fiscal policy analysis. An exception is the paper by Forni and Momigliano (2005). These authors estimate, for a panel of OECD countries, fiscal policy rules linking an indicator of discretionary fiscal policy, the cyclically-adjusted primary budget balance as percentage of potential GDP (hereafter $capb$), to the output gap and public debt. They use revised data for the policy instrument and the debt indicator, and revised versus real-time data for the output gap. They show that, when real-time information on cyclical conditions is incorporated, the discretionary stance of fiscal policy is gauged to be counter-cyclical, both in Euro Area and non-Euro Area OECD countries, but just during slowdowns.[7]
However, when the "intentional" stance of fiscal policy is considered, it might be of crucial importance to make correct assessments on the timeliness of information on the "fiscal instrument" itself. Typically, in fact, in each autumn of year $t-1$, fiscal authorities approve the budget for year t. Budget laws are designed on the basis of *ex-ante* projections on the state of the economy and on the perceived evolution of the public debt. In addition, the realization of planned fiscal measures depends importantly on implementations lags. Therefore, there might be relevant discrepancies between the discretionary fiscal measures as approved *ex-ante* and what observed several periods after decisions have been taken.[8]
This issue has been originally addressed in a former version of this paper (Cimadomo

[6]In the framework of monetary policy, since data on the potential output and output gaps (and to a minor extent the ones on inflation) are known with some accuracy only many quarters after the interest rate move has been decided; assessments based on monetary policy rules may be incorrect if revised data are used in the estimation.

[7]Loukoianova, Vahey and Wakerly (2003) construct a real-time data set for the U.S. primary surplus. However, they do not provide regression estimates based on their real-time data. Moreover, the fiscal policy indicator used is not cyclically adjusted. Therefore, they cannot discriminate between the effects of automatic stabilizers and discretionary measures.

[8]Note that this issue is not relevant for monetary policy analysis. In fact, central bankers can control movements in their operating instruments with great accuracy. In particular, short-term interest rates are subject to negligible revisions, and just for few days after the first release of data. The $capb$, on the contrary, is computed as difference between the nominal budget balance (net of interest payments) and the cyclical component of the balance, divided by potential output. Clearly, then, this indicator incorporates three sources of uncertainty and possible measurement errors: the level of nominal deficit, nominal output and potential output (which depends on estimates of the cyclical component of GDP). All of them are subject to considerable revisions.

(2006)), in which fiscal policy rules for a set of OECD countries are estimated using real-time data for the $capb$, and for all the other "ingredients" typically included in the rule.[9]
This paper complements Cimadomo (2006) and contributes to the literature on fiscal policy rules in the following directions:

- We formally show that in a panel regression framework, when both the dependent variable and the independent ones are contaminated by revision errors, and when these errors are correlated, the Fixed-Effects Least-Squares (FE-LS) estimator is inconsistent. The analytical form of the asymptotic bias is derived for simple fiscal rules, relating a fiscal policy indicator, here the $capb$, to the output gap as in Taylor (2000a) and Taylor (2000b).

- A real-time annual dataset is constructed by collecting data published in the December Issues of the OECD Economic Outlook from 1994 (Volume 56) to 2006 (Volume 80), for 19 OECD countries. Based on these data, it is shown that the bias incurred in estimating a simple fiscal rule using revised data, when the intentional fiscal policy stance is the object interest, can be accurately predicted based on the empirical correlations between revision errors in the $capb$ and in output gaps, and on other second-order moments. It is shown that the inclusion of real-time observations for the $capb$ revert the sign of the estimated parameter representing the cyclical sensitivity of discretionary fiscal policies, thereby indicating that the intentional fiscal policy stance seems to be counter-cyclical.

- More encompassing fiscal policy rules are estimated, where movements in the $capb$ are supposed to depend not only on cyclical conditions, but also on debt developments and on a set of other control variables, as in Galì and Perotti (2003). Again, it is documented that the use of revised observations for the fiscal policy instrument leads to an "attenuation bias", since the regression slope on the output gap is estimated to be lower, suggesting pro-cyclicality, than what obtained using real-time data. In particular it emerges that, *ex-ante*, fiscal policy is strongly counter-cyclical in expansions but substantially neutral during recessions.

[9] Golinelli and Momigliano (2006) include real-time observations of the primary balance on the right-hand side (RHS) of their regression equation, but they use revised data for dependent variable. More recently, Beetsma and Giuliodori (2007), to capture possible interdependances among fiscal policymakers in the European Union, "augment" a fiscal rule estimated for country i, where real-time values of the policy indicator are used, with an additional exogenous regressor representing a weighted average of the $capb$ over countries $j \neq i$.

- A two-stage procedure applied to the Hansen's (1999) threshold panel regression model is proposed to test the presence of non-linearities in the way discretionary fiscal policies respond to cyclical developments and to debt accumulations. It is found that the hypothesis of a switch in the *ex-ante* cyclical behavior of fiscal policy (from acyclicality or slight pro-cyclicality to counter-cyclicality) occurring when GDP is close to its equilibrium level is not rejected. Interestingly, the use of revised data does not allow to discriminate between any regime in the conduct of fiscal policy.

The rest of the paper is organized as follows. In Section 2 we present the reaction functions used in the evaluation of the fiscal policy stance, we document on the construction of the real-time dataset and we assess how revision errors in variables may affect estimation results; Section 3 is devoted to the presentation of the threshold regression model; Section 4 lays out the results; Section 5 presents some robustness exercises and Section 6 concludes.

2 Assessing the stance of fiscal policy in real-time

Attempts to model the behavior of fiscal authorities in terms of a "policy reaction function" are relatively recent in the empirical literature on fiscal policy. Taylor (2000*a*) and Taylor (2000*b*) argue that the conduct of fiscal policy may be well approximated by a rule (hereafter referred to as "simple" or "Taylor" fiscal rule) relating a measure of the fiscal policy stance to deviations of actual output from its equilibrium level, through a stable function $g(\cdot)$ as

$$fp_t = g(x_t) + \varepsilon_t, \tag{1}$$

where fp_t is the fiscal policy indicator, x_t is the output gap and ε_t are $i.i.d.$ residuals representing the "exogenous" or "unsystematic" component of the policy. Bohn (1998) suggests that sustainability issues may also play an important role in shaping the decisions of fiscal policymakers. In line with these arguments, the latest generation of fiscal rules incorporate the output gap but also the debt-output ratio (besides a set of additional controls) as explanatory variables in accounting for movements in the policy indicator, which is commonly selected to be the structural primary balance when the *discretionary* stance of fiscal policy is under investigation (see in particular Galì and Perotti (2003)).
Generally, "revised" data, i.e. observations from the latest available release, are used in the estimation of such rules. However, as suggested by Orphanides (2001), when the interest of the researcher is on the evaluation of the *intentional*, or *ex-ante*, fiscal

policy stance, all the information actually available to the policymaker at the time decisions have been taken should be used. In a fiscal policy reaction function framework, and in contrast with monetary Taylor rules, this information set should include real-time observations on the "operating instrument" in the hands of budgetary authorities, i.e. the discretionary component of the budget balance. In fact, a certain objective in terms of budget balance as planned in the current year may end up to be drastically different from what observed several years later, based on revised data. The potential impact of incorporating real-time information, in particular as concerns the fiscal policy indicator, on the assessment of the fiscal policy stance is explored in the following.

As spelled out more in detail in Section 2.1, the real-time data used in the analysis are drawn from the OECD Economic Outlook, and each vintage of data corresponds to the figures published in the December Editions from 1994 to 2006, for 19 OECD countries. This should allow to effectively capture the correct timing of the fiscal policy decision process, since budget laws for year $t + 1$ are passed at the end of the previous year and the December publication of the Economic Outlook plausibly reflects the information held by policymakers at the time of budgeting. In principle, growth projections and fiscal plans published by national statistical offices should be employed, since they should be more informative on the decisions of fiscal authorities. However, as documented by some authors (see e.g. Annett (2006) and Jonung and Larch (2004)), data released by national statistical agencies are often affected by a "political bias" inducing overly optimistic forecasts of GDP and the state of public finances. Hence, we rely on an independent institution such as the OECD as source of data, for it is less likely to be exposed to "political pressures" in compiling its statistics.

We proceed as follows. First, we formally document that the use of revised data in the estimation of a fiscal policy reaction function may yield biased results, when an *ex-ante* relation among variables is under investigation. Empirically, this is shown by estimating a simple fiscal rule "à la Taylor" where the fiscal policy indicator is the $capb$ and the explanatory variable is the (lagged) output gap. As a starting step, only revised data are employed in the regression. Then, real-time observations for the output gap are used. Finally, we provide regression estimates based on real-time figures for both the $capb$ and x.

Further, a more encompassing "backward-looking" specification of the fiscal policy reaction function, similar to the one proposed by Galì and Perotti (2003), is considered. Defining d the public debt (general government gross financial liabilities) to GDP ratio, a battery of four reaction functions (hereafter "baseline fiscal rules") is estimated, where the amount of real-time information incorporated is progressively

increased, from a "fully-revised" scenario to a "fully-real-time" one:
i) Revised-data: $capb, x, d$; no real-time data ("fully-revised" rule);

$$capb_{i,t} = \alpha_i + \rho capb_{i,t-1} + \beta x_{i,t-1} + \theta d_{i,t-1} + \psi\, emu_{i,t} + \varepsilon_{i,t}. \tag{2}$$

ii) Revised-data: $capb, d$; real-time data: x;

$$capb_{i,t} = \alpha_i + \rho capb_{i,t-1} + \beta x_{i,t-1|t-1} + \theta d_{i,t-1} + \psi\, emu_{i,t} + \varepsilon_{i,t}. \tag{3}$$

iii) Revised-data: $capb$; real-time data: d, x;

$$capb_{i,t} = \alpha_i + \rho capb_{i,t-1} + \beta x_{i,t-1|t-1} + \theta d_{i,t-1|t-1} + \psi\, emu_{i,t} + \varepsilon_{i,t}. \tag{4}$$

iv) No revised-data; real-time data: $capb, x, d$ ("fully-real-time" rule);

$$capb_{i,t|t-1} = \alpha_i + \rho capb_{i,t-1|t-1} + \beta x_{i,t-1|t-1} + \theta d_{i,t-1|t-1} + \psi\, emu_{i,t} + \varepsilon_{i,t}. \tag{5}$$

where $capb_{i,t|t-1}$ is the one-year-ahead forecast of the $capb$ as estimated in vintage $t-1$; the notation $z_{i,t-1|t-1}$ indicates the current-year estimate of z, for $z = capb, x, d$, provided in vintage $t-1$; the notation z_t denotes revised data, as conventional, and emu is a dummy variable which equals one from 1999 on for the countries having joined the European Monetary Union, and zero otherwise. The subscript i, with $i = 1, \ldots, N$, indexes the cross-section of countries.[10]
Moreover, for each of these models, two additional regressors are constructed by interacting x with a dummy variable which equals one when the output gap is positive (negative), and zero otherwise. Then, to capture possible asymmetries in the way fiscal policy reacts to the economic cycle, "conditional" fiscal rules including these two regressors are estimated. In the baseline exercises, FE-LS estimates are provided.[11]
The underlying assumption behind the proposed panel-regression models is that policymakers in the OECD countries behave uniformly, as far as reactions to output fluctuations and debt dynamics are concerned. Therefore, the common-across-countries

[10]Models (4) and (5) incorporate real-time values for the debt variable since, albeit some changes in accounting standards used to measure public liabilities have occurred over the considered period, $d_{t-1|t-1}$ is the level of the debt to GDP ratio *actually observed* by policymakers in period $t-1$. However, as it will be shown later, the use of revised data for d rather than real-time ones does not affect the results considerably.

[11]As well known, Least Squared estimators are asymptotically consistent for T large in dynamic panels (see Nickell (1981)). Moreover, compared to Instrumental Variables (IV) methods, results are not dependant on the choice of instruments. Nevertheless, as robustness checks, IV estimates are also shown in Section 5.

β and θ gauge, respectively, the "cyclical sensitivity" and the "sustainability concern" of fiscal authorities.[12] Possible (unobserved) country-specific heterogeneities are captured by the fixed-effect parameters α_i.

2.1 A real-time dataset for fiscal policy analysis

We construct a real-time annual dataset based on the December Issues of the OECD Economic Outlook from 1994 (Volume 56) to 2006 (Volume 80). The December Editions of the Outlook of each year t typically publish data spanning up to the previous sixteen years, "estimates" for the current-year and "forecasts" for years $t+1$ and $t+2$. The three reference indicators for which data have been collected are the output gap (deviation of actual GDP from potential GDP as percentage of potential GDP), the debt to GDP ratio (general government gross financial liabilities as percentage of nominal GDP) and the cyclically-adjusted primary balance as percentage of potential GDP.[13][14] Our real-time dataset is built on by inputting the current-year estimates, and the one year-ahead forecasts, of these variables.

The relationships between the real-time and revised observations of the variables used in the baseline estimations is defined as follows

[12]In this framework, a positive β indicates a "counter-cyclical" discretionary stance, since the *capb* increases during expansions (the so-called "saving for rainy days" policies) and drops during slowdowns. The policy stance is defined as "pro-cyclical" if β is negative, as discretionary fiscal policy decisions tend to exacerbate fluctuations in the economic cycle. In addition, the policy is characterized as "sustainable" when θ is positive and "unsustainable" when it is negative. In the former case, in fact, taxes are discretionarily increased and public spending reduced when debt dynamics are explosive. In the latter, discretionary policies contribute to worsen the state of public finance by increasing the debt-output ratio.

[13]The OECD began to release output gap data for all countries just in 1995. In 1994, however, estimates of the output gap were available for the G7 countries. For the remaining countries, the estimates provided by Forni and Momigliano (2005) are used. Note also that the OECD started to publish data on cyclically-adjusted *primary* balances just in 2002. Then, for the period 1994-2001, the *capb* has been constructed by adding net debt interest payments to the data on structural balances.

[14]Depending on the methodology employed to net out the effects of the economic cycle from fiscal aggregates, and on the elasticities used to gauge the "automatic" cyclical sensitivity of single budget items, estimated cyclically-adjusted indicators published by different institutions may be not equal (see Bouthevillain, Cour-Thimann, van den Dool, de Cos, Langenus, Mohr, Momigliano and Tujula (2001)). We use data from the OECD only for two main reason. First, the estimates provided by the IMF and the European Commission are broadly in line with the ones provided by the OECD and the potential impact on the estimation of fiscal rules is likely to be negligible. Second, the availability of several past issues of the OECD Economic Outlook makes the construction of a relatively large real-time dataset on fiscal variable feasible.

$$
\begin{aligned}
x_{i,t} &= x_{i,t|t} + \nu^{x}_{i,t},\\
capb_{i,t} &= capb_{i,t|t-1} + \nu^{capb}_{1,i,t},\\
capb_{i,t} &= capb_{i,t|t} + \nu^{capb}_{i,t},\\
d_{i,t} &= d_{i,t|t} + \nu^{d}_{i,t},
\end{aligned}
$$

where $\nu^{x}_{i,t}$, $\nu^{capb}_{i,t}$ and $\nu^{d}_{i,t}$ are the revision errors in the current-year estimates of the output gap, the $capb$ and d respectively; $\nu^{capb}_{1,i,t}$ is the revision error in the one-year-ahead forecast of the $capb$. Note that in the present framework, contrary to the standard approach, we consider as "correct" the observations reported in real-time, since we are interested in the *ex-ante* behavior of the policymaker.[15]

In addition, we also collect data on the current-year estimates of real GDP annual growth rates and of general government gross public debts as percentage of nominal GDP, according to the "Maastricht definition".[16] These two additional variables will be used in the proposed robustness exercises.

For all the indicators considered,we refer to revised data as the ones from the December 2006 Edition of the Economic Outlook.

The sample includes 19 OECD countries: Germany, Belgium, Austria, Finland, Spain, Greece, Ireland, Italy, France, Netherlands, Portugal, Sweden, Denmark, the United Kingdom, Norway, the United States, Canada, Japan and Australia.

Figure 1, 2 and 3 display three different vintages of data (1995,2000,2006) for the variables of interest. Even from a simple visual inspection, it can be noticed that the data from first two vintages are often largely different from what observed in the 2006 one. For instance, the Italian potential GDP for years 2000 and 2001, was perceived, in 2000, to be much stronger that it actually was, as shown by the 2006 estimates of the output gap for those years being around three percentage points higher than what published in the December 2000 Economic Outlook (Figure 1).

The mean absolute value of the revisions over the period of observation, as a summary statistic to gauge the magnitude of these measurement errors, is reported in Table 1. From the first column of Table 1 it emerges that, albeit for some countries (notably Belgium, the Netherlands and Australia) the output gap has been quite accurately measured, *on average*, over the last thirteen years; for the remaining ones

[15]Of course, this does not have any implication on the *absolute value* of revision errors and on their second-order moments.

[16]The Economic Outlook publishes data on government debt as defined by the Maastricht Treaty accounting rules just for the countries belonging to the European Union. Then, only data on fourteen out of the nineteen countries included in the original sample are available for this indicator.

revisions are generally large and amount to even more than two percentage points in the Finnish and Japanese case. Interestingly, Figure 2 shows that the $capb$ has also been inaccurately estimated for many of the countries in the sample. Given that this indicator is computed as a function of the output gap, countries for which x has been poorly measured also display large revision errors in the one-year-ahead and in the current-year forecasts of the $capb$ (Columns 2 and 3, Table 1). Moreover, the mean absolute value of revisions averaged over all the 19 countries shows that the $capb$ has been measured in real-time more poorly than the output gap. This depends on the fact that this fiscal indicator incorporates more sources of uncertainty: the nominal output and the potential one, as the output gap, but also the nominal deficit. The fourth column of the Table, and Figure 3, indicate that the level of debt-output ratio reported in year t is often remarkably different from what observed at the end of the sample, due to errors in measurement but also to possible changes in accounting rules. Revisions to this indicator are the largest for high debt countries such as Greece and Japan, but also for Norway.
Table 2 reports, for each country in the sample, the empirical correlations between revision errors in the fiscal policy instrument included in the "fully-revised" fiscal policy rule (2) and, in turn, revision errors in the RHS variables of that equation. Column 1 shows that $\nu_{1,i,t}^{capb}$ is negatively correlated with $\nu_{i,t-1}^{x}$ for all $i = 1, \ldots, 19$. Since the $capb$ is computed by subtracting a function of the output gap from the primary balance, it seems reasonable to observe that upward (downward) revisions in the output gap are associated with downward (upward) ones in the $capb$. Columns 2 indicates that the correlations between $\nu_{1,i,t}^{capb}$ and $\nu_{i,t-1}^{capb}$ are always positive and high whereas the ones between revision errors in the dependent variable and in the (lagged) debt are less uniform across countries, and approximately distributed around a zero mean value.

2.2 Why revision errors in the policy indicator may matter

As shown above, the main "ingredients" of the selected fiscal policy rules are contaminated by large measurement errors, which seem to be also highly cross-correlated. In the classical regression framework, the use of variables affected by measurement errors may invalidate the properties of commonly used estimators. A well known pitfall of Least Squares estimators, for instance, is that they become inconsistent when the *independent* variables included in the regression are measured with error (see for example Johnston and DiNardo (1997)). This eventually calls for IV methods, provided that appropriate instruments are found. Under standard assumptions, on the other hand, the presence of measurement errors in the *dependent* variable does not affect the consistency of LS estimators. This holds also in panel regressions.

Therefore, applying the FE-LS estimator to model (5) above, where the one-year-ahead forecast of the $capb$ is included on the LHS of the regression equation, should asymptotically (for $N, T \rightarrow \infty$) yield the same results as when revised data are used for the dependent variable.
Here we show that when the conventional assumption of uncorrelatedness between measurement errors is dropped, and this seems to be consistent with the results in Table 2, the FE-LS estimator becomes inconsistent not only due to measurement errors in the regressors, as in the standard case, but also to the fact that revision errors in the dependent variable and the independent ones might be correlated. The potential impact of correlated measurement errors has been already explored in the time series literature (see e.g. Haitovsky (1972)). In a panel regression framework, Biørn (1992) models the effects of applying the "Within", the "Between" and the difference transformation to the data, when observations on the regressors are contaminated by measurement errors. To our knowledge, the extension to the case of correlated measurement errors in the dependent variable and in the explanatory ones has not been formalized yet in the literature on panel data.
Let us consider a simple bivariate panel regression, or a Taylor fiscal rule according to the definition proposed above, where the dependent variable is $capb_{t|t-1}$ and the explanatory one is $x_{t-1|t-1}$. The structural equation of interest is the following

$$capb_{i,t|t-1} = \alpha_i + \beta x_{i,t-1|t-1} + \varepsilon_{i,t}, \tag{6}$$

where $\varepsilon_{i,t} \sim i.i.d(0, \sigma_\varepsilon^2)$. Suppose that the "true", or real-time, values for the $capb$ and x are not observed. Instead, we observe the "fallible", or revised, data denoted by $capb_{i,t}$ and x_{t-1}. As underlined before, in this framework we consider as "true" the real-time observations since we are interested in studying an *ex-ante* relation between variables.
Under a certain set of assumptions, and in particular allowing $\nu_{1,i,t}^{capb}$ and $\nu_{i,t-1}^{x}$ to be contemporaneously correlated, Appendix A formally shows that the asymptotic bias incurred in estimating (6) by FE-LS and using revised information is equal to

$$BIAS = \frac{1}{\sigma_{\tilde{x}^*}^2 + \sigma_{\tilde{\nu}^x}^2}(\sigma_{\tilde{\nu}^x \tilde{\nu}_1^{capb}} - \beta\sigma_{\tilde{\nu}^x}^2), \tag{7}$$

where $\sigma_{\tilde{x}^*}^2$ is the variance of the values of x, pooled across groups after removing individual means; $\sigma_{\tilde{\nu}^x}^2$ is the variance of the demeaned and pooled revision errors in x, $\sigma_{\tilde{\nu}^x \tilde{\nu}_1^{capb}}$ is the covariance between demeaned and pooled revisions errors in x and $capb$, and β is the true parameter. Equation (7) implies that we will tend to overestimate β when $\sigma_{\tilde{\nu}^x \tilde{\nu}_1^{capb}}$ is positive and to underestimate it when it is negative, i.e. an "attenuation bias" would arise. The standard textbook case assumes

$\sigma_{\tilde{\nu}^x \tilde{\nu}_1^{capb}} = 0$, and the bias will be influenced only by measurement errors in the independent variable. Based on the empirical second-order moments included in (7), it will be possible to accurately predict the size and the direction of the bias.[17]
When more complex specifications of the regression equation are considered, and in particular when the number of independent variables is large, the derivation of the analytical form of the bias is more cumbersome since it will depend on the cross-correlations among all measurement errors. In the case of the baseline fiscal rule (5) two additional regressors are added with respect to the simple specification (6): the lagged debt to GDP ratio and the lagged dependent variable. As in the simpler bivariate case, it can be shown in which direction the covariance between measurement errors in the dependent variable and in the output gap contributes to the overall bias.[18] To be noted, the estimator will be unbiased only in the (unlikely) case that all the second-order moments included in the functional form of the bias cancel out.

3 Testing for non-linearity in fiscal policy rules

This Section explores the possible presence of non-monotonic responses of fiscal authorities to the economic cycle and to public debt developments. The issue of whether fiscal policy behaves differently in various phases of the cycle has been extensively studied in the literature by relying on the notion of "good times" as periods of positive output gaps and "bad" ones as years, or quarters, in which the actual output is below the potential one (see, for example, Gavin and Perotti (1997), European Commission (2004) and OECD (2003)). The same approach is followed in Section 2, where the "conditional" fiscal rules proposed encompass positive and negative output gaps as separate regressors.
Some authors have however suggested that the "true" functional relation linking the fiscal policy indicator to the state of the economy might have an alternative form, suggesting that a switch in the policy behavior may occur around other phases of the cycle, for example when the output gap exceeds a certain threshold level. In particular, Manasse (2005) suggests that the cyclical sensitivity of fiscal policy may vary when slowdowns are particularly severe or upturns particularly strong, compared to intermediate states of the business cycle.[19]

[17] The use of IV estimators might contribute to mitigate this endogeneity problem, but unless a matrix of instruments, perfectly uncorrelated with measurement errors and residuals $\varepsilon_{i,t}$, and correlated with the revised $x_{i,t}$ is found, the estimator will still be inconsistent. In a time series regression framework, for instance, Orphanides (2001) estimates monetary Taylor rules using real-time data and IV methods and shows that the estimated coefficients are far from the ones obtained with revised data.

[18] Formal proof available from the author.

[19] In particular the author finds that, in a Barro-Gordon type of framework, and in the presence of

Threshold effects may be also at play as regards the response of fiscal policy to level of the government debt to GDP ratio. In the framework of the European Monetary Union and the Stability and Growth Pact, for instance, fiscal authorities may pursue more sustainable policies, attempting to reduce public debts, when the 60% ceiling is approached or exceeded. More in general, it can be expected that governments are more concerned about the sustainability of public finances when the public debt is high rather than when it is low. Hence, the relation linking the discretionary component of fiscal policy to debt developments might be non-monotonic, possibly switching around a certain threshold level of the debt indicator.

Based on Hansen's (1999) panel threshold model, and within a fiscal policy reaction function framework, here we test whether these threshold effects are statistically relevant. We will be able to discriminate between a single regime (linear fiscal rule), two regime (single threshold rule) and three regimes (double threshold rule) in the way the discretionary component of fiscal policy counteracts the economic cycle and government debt dynamics. Hansen's (1999) model is taken as a guideline as it does not rely on an ad hoc sample split of the data (as, for example, it is done when *positive* and *negative* output gaps are treated separately). Instead, it allows to *endogenously* identify, through a minimization criterion, single (or multiple) thresholds where a regime shift is more likely to occur (in a statistical sense). Furthermore, a bootstrap technique is proposed to assess the statistical significance of threshold effects, and an asymptotic theory (with fixed T as $N \to \infty$) is used to draw valid inference on parameters in different regimes.[20] Hansen's (1999) model has been developed for non-dynamic balanced panels. In the following we recall the key lines of that work, and we propose a two-stage procedure that allows applications to dynamic panel models. By employing revised and real-time data for 19 OECD countries, we explore the possible existence of non-linearities in both the *ex-post* and the *ex-ante*

limits on the deficit to GDP ratio, fiscal policy should be pro-cyclical during moderate economic downturns and counter-cyclical in more severe recessions. During mild slowdowns, in fact, governments are more likely to implement contractionary measures to avoid exceeding the deficit limit thereby triggering a further reduction in economic growth. During very "bad" economic times, on the contrary, the cost of abiding is too high and they find it optimal to brake the deficit rule and operate counter-cyclically through expansive policies.

[20]The choice of the methodological strategy used is also dictated by the particular type of non-linearity we are interested in (threshold effects) and by the panel structure of the dataset. A related application, but in a time series framework, is Favero and Monacelli (2005) which employs the Hamilton's (1989) regime switching model to the estimation of quarterly fiscal rules for the U.S. economy. They find that fiscal policy might be characterized as "active" (i.e. fiscal policymakers promote discretionary policies aiming at stabilizing output fluctuations) from the 1960s throughout the 1980s, then as "passive" (i.e. fiscal authorities are concerned about debt developments only) during the 1990s and "active" again since the start of the G. W. Bush Administration.

stance of fiscal policy.

3.1 Hansen's (1999) threshold panel regression model

This Section introduces the main building blocks of the panel threshold model by Hansen (1999). The single threshold case and the double threshold one will be analyzed, before proposing a two-stage procedure for applications in dynamic panels.

3.1.1 The single threshold model

Let y_{it}, z_{it} and q_{it} be the observed data from a balanced panel with $i = 1, \ldots, N$ and $t = 1, \ldots, T$. Defining $I(\cdot)$ the indicator function and μ_i the individual fixed effects, the (unobserved) structural equation linking the dependent variable y_{it} to the regressor z_{it} might be

$$y_{i,t} = \alpha_i + \beta z_{it} + \varepsilon_{i,t}, \tag{8}$$

or

$$y_{i,t} = \alpha_i + \beta_1 z_{it} I(q_{it} \leq \gamma) + \beta_2 z_{it} I(q_{it} > \gamma) + \varepsilon_{i,t}, \tag{9}$$

with $e_{i,t} \sim i.i.d(0, \sigma_e^2)$, depending on whether the relation between $y_{i,t}$ and $z_{i,t}$ changes when $q_{i,t}$ rises above a certain threshold γ. The procedure developed by Hansen (1999) allows to test the null hypothesis of no threshold effects

$$H_0 : \beta_1 = \beta_2.$$

Furthermore, if H_0 is rejected, inference on the threshold parameter is provided by testing

$$\tilde{H}_0 : \gamma = \gamma_0,$$

where γ_0 is the "true" threshold. Hansen's (1999) procedure for estimation and inference in single threshold models follows these main steps:

1. The observations on the threshold variable q_{it} are grouped across individuals and time, and sorted in ascending (or descending) order. From the $(NT \times 1)$ resulting **q** vector, select M distinct values $q_1 \ldots q_M$, after discarding the smallest and largest $\eta\%$, for some $\eta > 0$. These are the values used to search for $\hat{\gamma}$. For each q_j (or γ_j), perform the within transformation of equations

(9).[21] The demeaned equations are stacked and estimated by OLS. Defining $\tilde{\mathbf{e}}(\gamma)$ the vector of regression errors and $S_1(\gamma) = \tilde{\mathbf{e}}(\gamma)'\tilde{\mathbf{e}}(\gamma)$ the concentrated sum of squared errors; the threshold $\hat{\gamma}$ is estimated by minimizing $S_1(\gamma)$ over all values of γ. That is,

$$\hat{\gamma} = \operatorname*{argmin}_{\gamma} S_1(\gamma). \tag{10}$$

2. The H_0 hypothesis of no threshold effects is by the likelihood ratio statistic

$$F_1 = (S_0 - S_1(\hat{\gamma}))/\hat{\sigma}^2, \tag{11}$$

where S_0 is the sum of squared errors from the estimation of the linear model (8) and $\hat{\sigma}^2$ is the estimated variance of residuals from model (9). The asymptotic distribution of F_1 is non-standard, however a bootstrap procedure is proposed to derive asymptotically valid critical values.

3. When H_0 is rejected (i.e. when there is statistical evidence of a threshold effect), Hansen (2000) proves that $\hat{\gamma}$ is consistent for γ_0. The likelihood ratio statistic given by

$$LR_1(\gamma) = (S_1(\gamma) - S_1(\hat{\gamma}))/\hat{\sigma}^2, \tag{12}$$

is used to test $\tilde{H}_0 : \gamma = \gamma_0$. The likelihood ratio test is to reject for large value of $LR_1(\gamma_0)$. Theorem 1 in Hansen (1999) shows that under certain assumptions and $\tilde{H}_0 : \gamma = \gamma_0$,

$$LR_1(\gamma) \rightarrow_d \xi \tag{13}$$

as $n \rightarrow \infty$ where ξ is a random variable with distribution function

$$P(\xi \leq x) = (1 - exp(-x/2))^2. \tag{14}$$

The asymptotic distribution in (13) is pivotal, and it may be used to construct asymptotically valid confidence intervals. The distribution function (14) has inverse $c(\alpha) = -2\log(1 - \sqrt{1-\alpha})$, from which critical values can be calculated (for instance, the 1%, 5% and 32% critical values are 10.59, 7.35 and

[21] Hereafter γ is used for γ_j.

3.48 respectively). Finally, the "acceptance region" of confidence level $1-\alpha$ can be derived as that set of values for γ for which $LR_1(\gamma) \le c(\alpha)$. This can be visually seen by plotting $LR_1(\gamma)$ against a flat line at $c(\alpha)$.[22]

3.1.2 The double threshold model

The "true" model may incorporate more than one threshold. In Hansen's (1999) multiple threshold models, the procedure for estimation and inference on threshold parameters is more cumbersome, albeit intuitively similar, than what shown for the single threshold model. Here only the double threshold model is reviewed, since it is unlikely that the discretionary behavior of fiscal authorities might be characterized by more than three regimes. The double threshold regression model reads as

$$y_{i,t} = \alpha_i + \beta_1 z_{it} I(q_{it} \le \gamma_1) + \beta_2 z_{it} I(\gamma_1 < q_{it} \le \gamma_2) + \beta_3 z_{it} I(q_{it} > \gamma_2) + \varepsilon_{i,t}, \quad (15)$$

where $\gamma_2 > \gamma_1$. Estimation, testing for double threshold effects and confidence intervals constructions are performed as follows:

1. A sequential method is used to consistently estimate the γ_1 and γ_2 thresholds. First, estimate γ_1 as in step 1 of the single threshold model. A first-stage estimate $\hat{\gamma}_1$ is obtained. Next, fixing $\hat{\gamma}_1$, the second-stage threshold estimate is

   $$\hat{\gamma}_2^r = \underset{\gamma_2}{\text{argmin}}\, S_2^r(\gamma_2), \quad (16)$$

   where

   $$S_2^r(\gamma_2) = \begin{cases} S(\hat{\gamma}_1, \gamma_2) & \text{if} \quad \hat{\gamma}_1 < \gamma_2 \\ S(\gamma_2, \hat{\gamma}_1) & \text{if} \quad \gamma_2 < \hat{\gamma}_1 \end{cases} \quad (17)$$

   As shown in Bai (1997), $\hat{\gamma}_2^r$ is asymptotically efficient but the $\hat{\gamma}_1$ is not. Then, a third-stage estimator is proposed for the first threshold. This "refinement" estimate is

   $$\hat{\gamma}_1^r = \underset{\gamma_1}{\text{argmin}}\, S_1^r(\gamma_1) \quad (18)$$

[22] The distribution of the slope coefficient $\hat{\beta} = \hat{\beta}(\hat{\gamma})$ depends on the threshold estimate $\hat{\gamma}$. Hansen (2000) demonstrates that the dependence on the threshold estimate is of second-order importance. Therefore, $\hat{\beta}$ is asymptotically normal with covariance matrix estimated by $\hat{V} = (\mathbf{z}(\hat{\gamma})'\mathbf{z}(\hat{\gamma}))^{-1}\hat{\sigma}^2$, where $\mathbf{z}(\hat{\gamma})$ is the vector of stacked regressors, after removing individual means.

where

$$S_1^r(\gamma_1) = \begin{cases} S(\gamma_1, \hat{\gamma}_2) & \text{if} \quad \gamma_1 < \hat{\gamma}_2^r \\ S(\hat{\gamma}_2^r, \gamma_1) & \text{if} \quad \hat{\gamma}_2^r < \gamma_1 \end{cases} \tag{19}$$

2. To discriminate between one or two thresholds, and defining $S_2^r(\hat{\gamma}_2^r)$ and $\hat{\sigma}^2 = S_2^r(\hat{\gamma}_2^r)/n(T-1)$ the sum of squared errors and the estimated variance of second-stage residuals respectively; an approximate likelihood ratio test is proposed based on the statistic

$$F_2 = (S_1(\hat{\gamma}_1) - S_2^r(\hat{\gamma}_2^r))/\hat{\sigma}^2. \tag{20}$$

As before, the asymptotic distribution of F_2 is non-standard and Hansen (1999) develops a bootstrap procedure to construct appropriate critical values.

3. Finally, the $(1-\alpha)\%$ confidence intervals for γ_1 and γ_2 are derived based on

$$LR_2^r = (S_2^r(\gamma) - S_2^r(\hat{\gamma}_2^r))/\hat{\sigma}^2 \tag{21}$$

and

$$LR_1^r = (S_1^r(\gamma) - S_1^r(\hat{\gamma}_1^r))/\hat{\sigma}^2, \tag{22}$$

where $S_1^r(\gamma)$ and $S_2^r(\gamma)$ are defined in (17) and (19). The "no-rejection" regions are the set of values of γ such that $LR_1^r \leq c(\alpha)$ and $LR_2^r \leq c(\alpha)$.

3.2 A two-stage procedure applied to Hansen (1999)

The methodology proposed by Hansen (1999) has been developed for non-dynamic panel models and it cannot be automatically applied to dynamic ones. In a fiscal policy reaction function framework, however, a potential problem related to the inclusion of one (or more) lagged term of the dependent variable clearly arises. This holds for models (2), (3) and (4) above, where revised observations for the $capb$ are used. The panel regression model (5) is not properly dynamic, since the dependent variable is the one-year-ahead $capb$ forecast ($capb_{t|t-1}$) while the current-year-estimate of the $capb$ ($capb_{t-1|t-1}$) is included in the RHS of the equation. However, these two terms are likely to be highly correlated. Hence, we propose a two-stage procedure to address this problem.

In the first stage, the autoregressive coefficient ρ is estimated from the (linear) regressions (2), when only revised data are considered, and (5), when we are interested in the *ex-ante* behavior of fiscal policy. We label the first stage estimate as $\hat{\rho}_1$. In the second stage, $\hat{\rho}_1$ is treated as known and it is fixed in the following two non-linear panel regression models

$$\begin{aligned} capb_{i,t|t-1} &= \alpha_i + \hat{\rho}_1 capb_{i,t-1|t-1} + \beta_1 x_{i,t-1|t-1} I(x_{i,t-1|t-1} \leq \gamma_x) \\ &\quad + \beta_2 x_{i,t-1|t-1} I(x_{i,t-1|t-1} > \gamma_x) \\ &\quad + \theta d_{i,t-1|t-1} + \psi\, emu_{i,t} + \varepsilon_{i,t} \end{aligned} \tag{23}$$

and

$$\begin{aligned} capb_{i,t|t-1} &= \alpha_i + \hat{\rho}_1 capb_{i,t-1|t-1} + \beta x_{i,t-1|t-1} \\ &\quad + \theta_1 d_{i,t-1|t-1} I(d_{i,t-1|t-1} \leq \gamma_d) \\ &\quad + \theta_2 d_{i,t-1|t-1} I(d_{i,t-1|t-1} > \gamma_d) + \psi\, emu_{i,t} + \varepsilon_{i,t}, \end{aligned} \tag{24}$$

which are finally estimated based on Hansen (1999). Alternatively, and as a robustness check, $\hat{\rho}$ is imposed to be equal to one.[23]

As the two equations above show, possible non-linear reactions to cyclical developments and to the debt developments are modeled separately. First, we test whether the $capb$ may respond differently to the real activity conditional on the level of the output gap, and assuming that the reaction to the debt to GDP ratio is constant (model (23)). Secondly, and symmetrically, the fiscal policy indicator is allowed to react non-linearly to the debt indicator, keeping the sensitivity to the output gap invariant (model (24)). In this framework, γ_x and γ_d represent the (unknown) threshold levels associated with the output gap and the debt-GDP ratio. The notation in (23) and (24) refers to the "fully-real-time" and single threshold case. These models will be also estimated using revised data for all the variables included. Furthermore, the presence of three regimes (double threshold) will be tested.

[23]This also allows to avoid problems related to the construction of confidence intervals in the second stage, which should take into account the uncertainty stemming from the estimation of ρ from the first stage.

4 Results

4.1 Simple fiscal rules and bias prediction

Estimates of the simple fiscal rule (6), where the $capb$ is assumed to react only to the (lagged) output gap, and based on the whole panel of 19 OECD countries, are reported in Table 3. The estimated cyclical sensitivity parameter reported in the first Column indicates that when revised data are used for both variables, the stance of fiscal policy seems to be significantly pro-cyclical, as conventionally found in most of the literature on fiscal rules. When real-time values for the output gap are employed (Column 2), $\hat{\beta}$ is close to zero, and becomes insignificant. The third column displays the results obtained by using real-time data for both the dependent variable and the independent one. The estimated regression slope turns positive, indicating counter-cyclicality, and it is significant at the 1% level.[24]
Section 2.2 showed that applying the FE-LS estimator to revised data, when the "true" structural relation of interest is (6), yields inconsistent results if the dependent variable and the regressors are jointly contaminated by measurement errors, and if those errors are correlated. Table 1 documents that the one-year-ahead $capb$ and the current-year output gap have been rather inaccurately measured, as indicated by revision errors larger, on average, than one percentage point. Moreover, as shown in Table 2, the standard assumption of independence between revision errors seems to be strongly rejected in the data. By substituting the empirical counterparts of the variances and covariances of interest in the functional form of the bias (7), and by assuming that the "true" β is the one obtained from the real-time regression, we can accurately gauge the sign and the size of the bias incurred in using revised data. This is equal to

$$B\hat{I}AS = \frac{1}{\hat{\sigma}^2_{\tilde{x}^*} + \hat{\sigma}^2_{\tilde{\nu}^x}}(\hat{\sigma}_{\tilde{\nu}^x \tilde{\nu}_1^{capb}} - \beta \hat{\sigma}^2_{\tilde{\nu}^x}) = -0.52,$$

where $\hat{\sigma}^2_{\tilde{x}^*} = 1.35$, $\hat{\sigma}^2_{\tilde{\nu}^x} = 1.53$, $\hat{\sigma}_{\tilde{\nu}^x \tilde{\nu}_1^{capb}} = -1.01$ and $\beta = 0.33$. To be noted, by adding the estimated (negative) bias to the true β we get -0.19, a value very close to the regression slope obtained from revised data and equal to -0.14. This suggests that using revised data to assess the *ex-ante* stance of fiscal policy leads to an underestimation of the cyclical sensitivity coefficient, which becomes negative, (mistakenly)

[24]The R^2 of the Within regression is low across all the three experiments. This is due to the fact that the process governing the *capb* is very persistent. Hence, as shown below, the introduction of a term capturing inertia in budgetary decisions improves the regression fit dramatically. Note however that when real-time observations for both the variables included in the simple fiscal rule are used (Column 3), around 6% of the variability in the *capb* is explained by the output gap.

implying a pro-cyclical stance. As a byproduct, the relative contribution to the overall bias of the revision errors in x, and on the covariance between these revision errors and the ones in the $capb$, is computed. The former source accounts for 34% of the total bias, whereas the latter explain 66%, indicating that, in this framework, ignoring to model revision errors in the dependent variable may be drastically misleading.

4.2 Baseline fiscal rules

Table 4 reports the FE-LS estimates of the baseline fiscal rules proposed in Section 2, where movements in the discretionary fiscal policy indicator are supposed to depend not only on the output gap, but also on the debt-output ratio, on an autoregressive term capturing persistence in fiscal planning and on a dummy variable controlling for "EMU effects". In each of the four models considered, from equation (2) to equation (5), the output gap is first included "unconditionally". Then, negative and positive output gaps are incorporated as separate regressors, and the associated regression slopes are β_1 and β_2 respectively.
Column 1 and 2 display the results when only ex-post data are used in the regressions. The cyclical sensitivity parameter estimate is -0.13 and it is highly significant, pointing to pro-cyclicality in the *ex-post* fiscal policy stance, consistently with most of the literature using revised data. This holds in particular during economic upturns, whereas the low and insignificant estimate for β_1 signals an acyclical stance during downturns. When real-time values of the output gap are employed, $\hat{\beta}$ becomes insignificantly different from zero, suggesting acyclicality, whereas $\hat{\theta}$ remains positive but it is more precisely estimated (Column 3 and 4).
The introduction of real-time information on the debt indicator does not alter the picture much (Columns 5 and 6), but when the one-year-ahead forecasts of the $capb$ and the current-year estimates of it are used in the regression, the results radically change. In particular, $\hat{\beta}$ reverts its sign and becomes significantly (at the 10% level) positive, signalling counter-cyclicality, as shown in Column 7. When upswings and slowdowns in the economic cycle are analyzed separately (Column 8), it emerges that the counter-cyclical reaction of fiscal policy is particularly strong when the actual output is above its equilibrium level, while the fiscal policy stance seems neutral in the opposite case. In particular, during upturns, a one percentage point increase in the output gap induces a 0.29 percentage point fiscal tightening. This contrasts with Forni and Momigliano (2005) who, using real-time data for the output gap, and revised data for the policy indicator and all the other variables, find that the stance of fiscal is counter-cyclical, but just during recessions. Interestingly, the *ex-ante* behavior of fiscal authorities does not appear to be characterized by a "sustainability concern" related to the level of public indebtedness, as indicated by an estimate of

θ not statistically different from zero. This result is at odds with Bohn (1998) who, based on U.S. revised data, finds that the primary balance is an increasing function of the debt-output ratio thus implying a sustainable conduct of fiscal policy in that country. The R^2 of the "fully-real-time" regression is remarkably higher compared to ones of the previous models thereby indicating the appropriateness of this specification when the interest is on the *intentional* fiscal policy stance. Finally, the negative sign of $\hat{\psi}$ across all the experiments shows that the conduct of fiscal policy has been more lax since 1999 than before in the countries that joined the single currency. This is consistent with the mainstream view associating the strongest fiscal consolidation efforts with the years preceding the start of the EMU.

4.3 Threshold effects in fiscal policy

The Hansen's (1999) panel-threshold model and the two-stage refinement proposed above are applied to study the (possibly non-linear) nature of the relationship linking the discretionary component of fiscal policy, as represented by the $capb$, to the output gap and government debt. By employing real-time data, we will be able to assess whether the *ex-ante*, or *intentional*, conduct of fiscal policy has been characterized by a single regime, or whether it has changed around certain (at most two) threshold levels of our indicators for the real activity and public indebtedness. Since, as documented before, any assessment on the stance of fiscal policy seems to heavily depend on whether *ex-ante* or *ex-post* data are used, we will also test for possible threshold effects in fiscal rules estimated on revised data.

The implementation of the testing procedure, based on the 19 OECD country dataset spanning the period 1994 to 2006, follows the steps laid out in Section 3.1 and 3.2.[25] Table 5 reports the F_1 and F_2 statistics, along with their p-values constructed from 1000 bootstrap replications, used to test whether the reaction of the $capb$ to the output gap is constant over the business cycle or varies in different phases of it. The results in Column 1 and 2, based on real-time data, suggest that the null of no threshold effects is rejected at the 5% level both when the autoregressive coefficient from the first-stage estimation is 0.78 and when $\hat{\rho}_1$ is set equal to 1. The presence of three regimes is however rejected at the 10% level, as indicated by the p-values associated with the F_2 statistics. This suggests that a single threshold fiscal reaction function (as equation (23)) should be the appropriate model for analyzing the cyclical stance of fiscal policy. Figure 4 shows that the point estimate of the threshold γ_x in the single threshold model is -1.2 and corresponds to the value at which the likelihood

[25] For each experiment performed, the trimming parameter η is fixed to 20 such that we can ensure that at least 45 observations ($\simeq N \times T \times \eta/100$) lie in each regime (see step 1 in Section 3.1.1).

ratio LR_1 (see equation (12)) hits the zero axis. The $\hat{\beta}_1$ and $\hat{\beta}_2$ regression estimates reported in Table 6 suggest that the fiscal policy stance seems neutral when the output gap is below this threshold, whereas strongly and significantly counter-cyclical when it is above it.[26] The 68% and 95% confidence intervals for γ_x displayed in Table 6, and respectively constructed from the values of the threshold variable for which the LR_1 lies below the solid green line and the dotted green one in Figure 4, are quite large and include values of x in the zero region together with "moderate" expansions and recessions.

These findings indicate that asymmetric effects seem to be at play when the *intentional* behavior of fiscal policy is considered, and that it is reasonable to model a switch in the policy stance as occurring when the actual output is close to the potential one. By contrast, when the estimation is based on revised data, the low values of the F_1 statistic, reported in Column 3 and 4 of Table 5, imply that we cannot discriminate between different regimes in the behavior of fiscal policy, if analyzed *ex-post*.

The results on the possible presence of non-linear effects in the discretionary fiscal policy response to movements in government debt, as measured by the gross financial liabilities to GDP ratio, are reported in the top panel of Table 7. The estimated F_1 statistics suggest that threshold effects have always to be rejected, no matter whether real-time or revised data are used in the estimation. Further, since values of the general government public debt based on the notion of gross financial liabilities may partially differ from the ones based on the "Maastricht definition" of it, we also test for threshold effects when this latter indicator is used. Only data on eleven European Union countries are employed, since the OECD does not publish any data on this indicator for the five OECD non-EU countries included in the original sample of 19 countries.[27] The results presented in the bottom panel of Table 7 suggest that the hypothesis of multiple regimes is again rejected, both in the real-time and the revised case. It is however interesting to note that, as the bottom graph of Figure 4 shows for the $\hat{\rho}_1 = 1$ and "fully-real-time" case, the likelihood ratio LR_1 is minimized when the threshold variable is at 80.5%. When this threshold is used in the estimation, the $\hat{\theta}_1$ and $\hat{\theta}_2$ "regime-dependent" regression slopes, representing the discretionary response of fiscal policy to debt when this is below (above) the threshold level (see equation

[26] The double threshold model, albeit rejected by the data, indicates that a second threshold might be located at -3.2. When the output gap is below this threshold, the estimated cyclical sensitivity points to counter-cyclicality. This finding is consistent with Manasse's (2005) model which predicts a counter-cyclical fiscal policy stance during "severe recessions", in particular for countries adopting the Stability and Growth Pact.

[27] Austria, Finland and Sweden are dropped from the *balanced* panel since the December 1994 OECD Economic Outlook does not provide data on these countries for the same year.

(24)), are equal to 0.0182 and 0.0039 respectively. The former is significant at the 95% level while the latter is statistically insignificant. This would imply that, over the last thirteen years, European Union governments have reacted in a sustainable way to the accumulation of public debt when its level was relatively low. Contrary to what it might be expected, however, these estimates would also indicate that the reaction has been weaker when the 60% limit has been (largely) exceeded, suggesting that the "dissuasive arm" of the Stability and Growth Pact failed to encourage more virtuous policies when the level of public debt was particularly high.

5 Robustness checks

Table 8 presents the results from some robustness exercises. The benchmark estimates of the "fully-real-time" fiscal rule (from Columns 7 and 8 of Table 4) are reported in Columns 1 and 2.
In the first experiment, since estimates of the output gap depend upon the specific methodology employed to compute the potential output, and given that the OECD follows a "production function" approach (see Giorno, Richardson, Roseveare and van den Noord (1995)), we propose an alternative real-time measure of the output gap based on the Hodrick-Prescott filter. For each country i and each vintage t (from 1994 to 2006), we reconstruct real GDP series, *in levels*, and we compute the output gap by following these steps:[28]

1. The available observations on real GDP growth rates from $t_1 = t - T$ to t, are collected;[29]

2. We normalize at 100 the first value of real GDP in levels, corresponding to year $t_0 = t_1 - 1$. All the remaining observations are computed by recursively applying the annual GDP growth rates;

3. A trend in the (reconstructed) GDP series in levels is estimated by the Hodrick-Prescott filter, with a smoothing parameter set equal to 100;

4. The new real-time output gap for year t, defined as $x^{hp}_{i,t|t}$, is computed as deviation of GDP from the value of the trend estimated for the same year. Series of positive (negative) output gaps are constructed by interacting $x^{hp}_{i,t|t}$ with a dummy variable equal to one when $x^{hp}_{i,t|t} > 0$ ($x^{hp}_{i,t|t} \leq 0$), and zero otherwise.

[28] Note that the OECD Economic Outlook does not report data on real GDP *in levels*, but just in terms of growth rates.

[29] T is set equal to 13 since this is the maximum, and common across countries and vintages, horizon over which data on real GDP growth are reported in the Economic Outlook.

The results laid out in Column 3 show that, when this real-time measure of the output gap is used in the regressions, the estimated coefficient representing the cyclical sensitivity of fiscal policy is 0.10, and significant at the 99% level. When upswings and slowdowns are considered separately, the estimated slopes point to counter-cyclicality during buoyant economic times, whereas the policy stance appears to be neutral when output is below the estimated long-run level.

Next, we use an different indicator of cyclical conditions, represented by real GDP growth rates (as percentage change from previous years), measured in real-time, replacing the output gap. The underlying idea is that policymakers might not be able to compute the potential output, or might not want to rely on such an uncertain indicator in designing their policies, and may respond only to the output growth as a measure of real activity. The positive and 95% significant $\hat{\beta}$ in Column 5 points to counter-cyclicality, as in the benchmark case. Then, positive and negative "growth gaps" are included as separate exogenous variables. These regressors are constructed by removing individual means from the real-time GDP growth rate series for each country i. Positive (negative) growth gaps are derived by multiplying the demeaned series, named $gdp_{i,t|t}$, by an indicator function which takes value one (zero) when $gdp_{i,t|t} > 0$, and zero (one) otherwise. It emerges that the fiscal policy stance seems counter-cyclical when growth is above its average, whereas acyclical in the opposite case.

Furthermore, a "forward looking" specification of the fiscal rule is estimated, where the one-year-ahead forecast of the output gap (as published by the OECD) is included as measure of real activity (Column 7 and 8). This is consistent with the possibility that fiscal policy authorities may react to *expected* cyclical conditions, rather than *current* ones. In this case, an endogeneity bias in estimation may occur, stemming from a possible inverse causality between the $capb$ and the cyclical indicator. Hence, regression estimates are based on an IV approach where the instruments used are the current-year estimate of the GDP growth rate for year $t-1$ ($gdp_{i,t-1|t-1}$) and the $t-1$ current-year output gap (unweighted) averages over all the OECD countries considered (excluding country i). The Sargan test suggests that the over-identifying restrictions induced by the proposed instruments are valid, both in "unconditional" case and in the "conditional" one. The estimates shown in Column 7 indicate that the unconditional reaction to cyclical fluctuations is counter-cyclical and statistically significant. Moreover, fiscal authorities seem to respond very asymmetrically to *expected* upturns and slowdowns in the economic cycle as indicated by a $\hat{\beta}_2$ coefficient equal to 0.46, and significant at a 99% level, and a $\hat{\beta}_1$ slope close to zero and insignificant (Column 8).

Finally, we control for the possibility that the "political cycle" may play an important role in shaping the behavior of fiscal authorities, as in Buti and van den Noord (2004). The benchmark regression (equation (5)) is augmented by a dummy variable, named $elect_{i,t}$, taking value of one in parliamentary election years and zero otherwise.[30] As Column 9 and 10 show, the coefficient associated with this regressor is significant and of the expected negative sign indicating that the occurrence of an election leads to more fiscal profligacy, thereby reducing public savings. The sign and the size of all the other coefficients are however not importantly affected suggesting that the inclusion of this additional exogenous variable is not relevant in the assessment of the cyclical behavior of fiscal policy.

6 Conclusions

This paper has shown that, in fiscal policy analysis, realistic assumptions about the timeliness at which information is available to policymakers are of key importance. When the object of interest is the *intentional* stance of fiscal policy, real-time observations of the operating instrument used by fiscal policymakers should be employed in the estimation of fiscal rules. We demonstrate that the sign and the size of the bias incurred in estimating a fiscal rule on revised data can be accurately predicted based on empirical correlations among measurement errors, and on other second-order moments. In particular, our findings suggest that the use of updated observations would mistakenly point to a pro-cyclical fiscal policy stance in industrialized countries over the 1994-2006 period, whereas real-time data indicate the contrary, at least as long as economic expansions are concerned.

Further, formal tests based on Hansen (1999) are performed to explore whether the discretionary behavior of fiscal policy might have been characterized by multiple regimes. It emerges that a switch in the intentional fiscal policy stance, from neutral (or slightly pro-cyclical) to counter-cyclical, is likely to occur when output is around its equilibrium level. On the other hand, we find that the use of revised data does not allow to identify any significant threshold effect in the cyclical conduct of fiscal policy. Threshold effects are always rejected as regards the response of fiscal policy to debt accumulation, both when real-time or revised data are used in the estimation. Overall, these findings cast some doubts on the efficacy of discretionary fiscal policies to fine tune the business cycle. In fact, albeit the *intentional* stance of the policy seems genuinely counter-cyclical, *ex-post* we find a pro-cyclical behavior. This suggests that the long and uncertain lags behind the budgetary process, coupled with difficulties in

[30] Data on election years are taken from the website of the International Institute for Democracy and Electoral Assistance (http://www.idea.int/vt/parl.cfm) and from the Election Resources on the Internet website (http://electionresources/org).

correctly measuring the output gap at the time of budgeting, have probably prevented stabilizing fiscal measures to be timely implemented over the economic cycle.

References

Annett, A. (2006), Enforcement and the Stability and Growth Pact: How Fiscal Policy Did and Did Not Change Under Europe's Fiscal Framework, IMF Working Paper No. 06/116.

Bai, J. (1997), 'Estimating Multiple Breaks One at a Time', *Econometric Theory* **13**, 315–352.

Beetsma, R. and Giuliodori, M. (2007), On the Relationship between Fiscal Plans in the European Union: An Empirical Analysis Based on Real-Time Data, CEPR Discussion Paper No. 6088.

Biørn, E. (1992), 'The Bias of Some Estimators for Panel Data Models with Measurement Errors', *Empirical Economics* **17**, 51–66.

Bohn, H. (1998), 'The Behavior of U.S. Public Debt and Deficits', *The Quarterly Journal of Economics* **113**(3), 949–963.

Boivin, J. (2005), Has US Monetary Policy Changed? Evidence from Drifting Coefficients and Real-Time Data, NBER Working Paper No. 11314.

Bouthevillain, C., Cour-Thimann, P., van den Dool, G., de Cos, P. H., Langenus, G., Mohr, M., Momigliano, S. and Tujula, M. (2001), Cyclically Adjusted Budget balances: An Alternative Approach, ECB Working Paper No. 77.

Brunila, A., Buti, M. and in't Veld, J. (2002), Fiscal Policy in Europe: How Effective are Automatic Stabilisers?, Economic Paper No. 177, European Commission.

Buti, M. and Sapir, A. (2006), Fiscal Policy in Europe: The Past and the Future of EMU Rules from the Perspective of Musgrave and Buchanan, CEPR Discussion Paper No. 5830.

Buti, M. and van den Noord, P. (2004), 'Fiscal Discretion and Election in the Early Years of the EMU', *Journal of Common Market Studies* **42**(4), 737–756.

Chari, V. and Kehoe, P. J. (1999), Optimal Fiscal and Monetary Policy, *in* J. B. Taylor and M. Woodford, eds, 'Handbook of Macroeconomics', Vol. 1 of *Handbook of Macroeconomics*, Elsevier, chapter 26, pp. 1671–1745.

Cimadomo, J. (2006), Testing Non-Linearity in Fiscal Policy: New Evidence Based on Real-Time Data. CEPII and ECARES-Free University of Brussels, *mimeo*.

Croushore, D. and Stark, T. (2001), 'A Real-Time Dataset for Macroeconomists', *Journal of Econometrics* **105**, 111–130.

European Commission (2004), The Pro-Cyclicality of Fiscal Policy in EMU, *in* 'Quarterly Report on the Euro Area', Vol. 3.

Favero, C. and Monacelli, T. (2005), Fiscal Policy Rules and Regime (In)Stability: Evidence from the U.S, Working Paper No. 282, IGIER (Innocenzo Gasparini Institute for Economic Research), Bocconi University.

Forni, L. and Momigliano, S. (2005), 'Cyclical Sensitivity Of Fiscal Policies Based On Real-Time Data', *Applied Economics Querterly* **50**(3), 299–326.

Galì, J. and Perotti, R. (2003), 'Fiscal Policy and Monetary Integration in Europe', *Economic Policy* **18**(37), 533–572.

Gavin, M. and Perotti, R. (1997), Fiscal Policy in Latin America, *in* Ben S. Bernanke and Julio Rotemberg, ed., 'NBER Macroeconomic Annual', The MIT Press, chapter 1.

Giannone, D., Reichlin, L. and Sala, L. (2005), Monetary Policy in Real Time, CEPR Discussion Paper No. 4981.

Giorno, C., Richardson, P., Roseveare, D. and van den Noord, P. (1995), Estimating Potential Output, Output Gaps and Structural Budget balances, OECD Economics Department Working Paper No. 152.

Golinelli, R. and Momigliano, S. (2006), 'Real-Time Determinants of Fiscal Policies in the Euro Area', *Journal of Policy Modeling* **28**(9), 943–964.

Haitovsky, Y. (1972), 'On Errors of Measurement in Regression Analysis in Economics', *International Statistical Review* **40**(1), 23–35.

Hamilton, J. D. (1989), 'A New Approach to the Economic Analysis of Nonstationary Time Series and the Business Cycle', *Econometrica* **2**(57), 357–384.

Hansen, B. E. (1999), 'Threshold Effects in Non-Dynamic Panels: Estimation, Testing, and Inference', *Journal of Econometrics* **93**(2), 345–368.

Hansen, B. E. (2000), 'Sample Splitting and Threshold Estimation', *Econometrica* **68**(3), 575–604.

Ironside, B. and Tetlow, R. J. (2005), Real-Time Model Uncertainty in the United States: the Fed from 1996-2003, CEPR Discussion Paper No. 5305.

Johnston, J. and DiNardo, J. (1997), *Econometric Methods*, McGraw-Hill International Editions.

Jonung, L. and Larch, M. (2004), Improving Fiscal Policy in the EU: The Case for Independent Forecasts, European Commission Economics Paper No. 210.

Kaminsky, G. L., Reinhart, C. M. and Vegh, C. A. (2004), When it Rains, it Pours: Procyclical Capital Flows and Macroeconomic Policies, NBER Working Paper No. 10780.

Lane, P. R. (2003), 'Business Cycles and Macroeconomic Policy in Emerging Market Economies', *International Finance* **6**(1), 89–108.

Loukoianova, E., Vahey, S. P. and Wakerly, E. C. (2003), A Real Time Tax Smoothing Based Fiscal Policy Rule, Computing in Economics and Finance No. 118, Society for Computational Economics.

Lucas, R. E. and Sargent, T. J. (1978), After Keynesian Macroeconomics, *in* 'After the Phillips Curve: Persistence of High Inflation and High Unemployment', Federal Reserve Bank of Boston.

Manasse, P. (2005), Deficit Limits, Budget Rules and Fiscal Policy, IMF Working Paper No. 120, International Monetary Fund.

Nickell, S. (1981), 'Biases in Dynamic Models with Fixed Effects', *Econometrica* **49**, 1417–1426.

OECD (2003), Economic Outlook, No. 74 (December).

Orphanides, A. (2001), 'Monetary Policy Rules Based on Real-Time Data', *American Economic Review* **91**(4), 964–985.

Sargent, T. J. and Wallace, N. (1975), '"Rationa" Expectations, the Optimal Monetary Instrument, and the Optimal Money Supply Rule', *Journal of Political Economy* **83**(2), 241–54.

Taylor, J. (2000*a*), 'Reassessing Discretionary Fiscal Policy', *Journal of Economic Perspectives* **14**(3), 21–36.

Taylor, J. (2000*b*), The Policy Rule Mix: A Macroeconomic Policy Evaluation, *in* Guillermo Calvo, Rudiger Dornbusch, and Maurice Obstfeld, ed., 'Money, Capital Mobility and Trade, Essays in Honor of Robert Mundell', the MIT Press.

A Appendix

This appendix shows that in a simple bivariate panel regression framework the FE-LS estimator is inconsistent, when both the dependent variable y and the regressor z are contaminated by measurement errors, and under the condition that these measurement errors are cross-correlated. The analytical form of the (asymptotic) bias is derived.

Let the scalars $y_{i,t|t}$ and $z_{i,t|t}$ denote the "true" values for the variable y and x. Clearly, in this setup, the notation used suggests that the true values correspond to the observations available in "real-time". The structural equation of interest is the following

$$y_{i,t|t} = \alpha_i + \beta z_{i,t|t} + \varepsilon_{i,t}, \qquad i = 1, \ldots, N; t = 1, \ldots, T, \tag{A.1}$$

where $\varepsilon_{i,t} \sim i.i.d(0, \sigma_\varepsilon^2)$. Suppose that the variables $y_{i,t|t}$ and $z_{i,t|t}$ are not actually observed. Instead, the "revised" data $y_{i,t}$ and $z_{i,t}$ are observed and used to estimate β, where

$$y_{i,t} = y_{i,t|t} + \nu_{i,t}^y, \tag{A.2}$$

$$z_{i,t} = z_{i,t|t} + \nu_{i,t}^z, \tag{A.3}$$

and where $\nu_{i,t}^y$ and $\nu_{i,t}^y$ are non-autocorrelated measurement errors in y and z.[31]Let this set of assumptions hold

$$Cov(y_{i,t|t}, \nu_{i,t}^y) = 0, \tag{A.4}$$

$$Cov(z_{i,t|t}, \nu_{i,t}^z) = 0, \tag{A.5}$$

$$Cov(y_{i,t|t}, \nu_{j,t}^y) = 0, \text{ for } i \neq j; \; i, j \in [1, N], \tag{A.6}$$

$$Cov(z_{i,t|t}, \nu_{j,t}^z) = 0, \text{ for } i \neq j; \; i, j \in [1, N]. \tag{A.7}$$

Within each group i, the measurement errors are supposed to follow a generic distribution F

$$\begin{pmatrix} \nu_{i,t}^y \\ \nu_{i,t}^z \end{pmatrix} \sim F\left(\begin{bmatrix} \mu_i^{\nu^y} \\ \mu_i^{\nu^z} \end{bmatrix}, \begin{bmatrix} \sigma_{\nu^y}^2 & \sigma_{\nu^y \nu^z} \\ \sigma_{\nu^z \nu^y} & \sigma_{\nu^z}^2 \end{bmatrix} \right), \tag{A.8}$$

[31]To be noted, in the present framework the relationship between the correct values and "fallible" observations is reversed compared to the conventional approach, since here the interest is on real-time data.

with non-zero (possibly different across groups) means and contemporaneous cross-covariances assumed to be different from zero ($\sigma_{\nu^z \nu^y} \neq 0$). Furthermore, the measurement errors in y and z are allowed to be correlated across groups: $Cov(\nu^{y_i}, \nu^{z_j}) = \sigma_{\nu^{y_i} \nu^{z_j}} = \sigma_{\nu^y \nu^z}, \forall i, j \in [1, N]$.
Rearranging (A.2) and (A.3) as $y_{i,t|t} = y_{i,t} - \nu^y_{i,t}$ and $z_{i,t|t} = z_{i,t} - \nu^z_{i,t}$ and substituting these expressions into (A.1) we obtain

$$y_{i,t} - \nu^y_{i,t} = \alpha_i + \beta(z_{i,t} - \nu^z_{i,t}) + \varepsilon_{i,t}. \tag{A.9}$$

The within transformation is performed on the set of equations (A.9). Define $\bar{w}_i = \frac{1}{T}\sum_t w_{i,t}$ for $w = y, z, \nu^y, \nu^z$; from (A.9) we get

$$\bar{y}_i - \bar{\nu}^y_i = \alpha_i + \beta(\bar{z}_i - \bar{\nu}^z_i) + \varepsilon_{i,t} - \bar{\varepsilon}_i. \tag{A.10}$$

Subtracting (A.10) from (A.9), and recalling that $\varepsilon_{i,t}$ has zero mean, gives

$$(y_{i,t} - \bar{y}_i) - (\nu^y_{i,t} - \bar{\nu}^y_i) = \beta[(z_{i,t} - \bar{z}_i) - (\nu^z_{i,t} - \bar{\nu}^z_i)] + \varepsilon_{i,t}. \tag{A.11}$$

These NT equations can be expressed as follows

$$\tilde{y}_{i,t} = \beta\tilde{z}_{i,t} + \tilde{\nu}^y_{i,t} - \beta\tilde{\nu}^z_{i,t} - \varepsilon_{i,t}, \tag{A.12}$$

where the notation $\tilde{w}_{i,t} = w_{i,t} - \bar{w}_i$ (for $w = y, z, \nu^y, \nu^z$) denotes demeaned values. The FE-LS estimator of β is obtained by pooling across the groups i the demeaned equations in (A.12) and applying ordinary least squares.
Stacking the equations in (A.12) we get

$$\tilde{\mathbf{y}} = \beta\tilde{\mathbf{z}} + \tilde{\mathbf{v}}^y - \beta\tilde{\mathbf{v}}^z + \mathbf{e}, \tag{A.13}$$

where $\tilde{\mathbf{y}}$,$\tilde{\mathbf{z}}$,$\tilde{\mathbf{v}}^y$,$\tilde{\mathbf{v}}^z$ and $\mathbf{e}$ are $(NT \times 1)$ column vectors. Defining $\tilde{\mathbf{y}}^*$ and $\tilde{\mathbf{z}}^*$ the $(NT \times 1)$ column vectors obtained by stacking the (demeaned) true values for y and x, from (A.2) and (A.3) it follows that $\tilde{\mathbf{y}} = \tilde{\mathbf{y}}^* + \tilde{\mathbf{v}}^y$ and $\tilde{\mathbf{z}} = \tilde{\mathbf{z}}^* + \tilde{\mathbf{v}}^z$. Indicating with $\tilde{y}_\tau$,$\tilde{z}_\tau$,$\tilde{y}^*_\tau$,$\tilde{z}^*_\tau$,$\tilde{\nu}^y_\tau$,$\tilde{\nu}^z_\tau$, with $\tau = 1, \ldots, NT$, the scalars from the vectors $\tilde{\mathbf{y}}$,$\tilde{\mathbf{z}}$,$\tilde{\mathbf{y}}^*$,$\tilde{\mathbf{z}}^*$,$\tilde{\mathbf{v}}^y$,$\tilde{\mathbf{v}}^z$ we may write (subscripts are dropped for simplicity)

$$\tilde{y} = \tilde{y}^* + \tilde{\nu}^y, \tag{A.14}$$
$$\tilde{z} = \tilde{z}^* + \tilde{\nu}^z. \tag{A.15}$$

From (A.4) to (A.7) it follows that $Cov(\tilde{y}^*, \tilde{\nu}^y) = 0$ and $Cov(\tilde{z}^*, \tilde{\nu}^z) = 0$; from (A.8) we have

$$\begin{pmatrix} \tilde{\nu}^y \\ \tilde{\nu}^z \end{pmatrix} \sim F\left(\begin{bmatrix} 0 \\ 0 \end{bmatrix}, \begin{bmatrix} \sigma^2_{\tilde{\nu}^y} & \sigma_{\tilde{\nu}^y \tilde{\nu}^z} \\ \sigma_{\tilde{\nu}^z \tilde{\nu}^y} & \sigma^2_{\tilde{\nu}^z} \end{bmatrix} \right), \tag{A.16}$$

By assumption, second-order moments of the variables of interest exist. It then follows that

$$plim(\frac{1}{NT}\tilde{\mathbf{z}}'\tilde{\mathbf{v}}^y) = \sigma_{\tilde{\nu}^y \tilde{\nu}^z},$$
$$plim(\frac{1}{NT}\tilde{\mathbf{z}}^{*\prime}\tilde{\mathbf{z}}^*) = \sigma^2_{\tilde{z}^*},$$
$$plim(\frac{1}{NT}\tilde{\mathbf{z}}'\tilde{\mathbf{z}}) = \sigma^2_{\tilde{z}^*} + \sigma^2_{\tilde{\nu}^z},$$
$$plim(\frac{1}{NT}\tilde{\mathbf{z}}'\tilde{\mathbf{v}}^z) = \sigma^2_{\tilde{\nu}^z},$$

where $plim$ is the probability limit, for $N \to \infty$ and $T \to \infty$. From (A.13), and under the assumption $plim(\frac{1}{NT}\tilde{\mathbf{z}}'\mathbf{e}) = 0$, the FE-LS estimator of β is obtained as

$$\hat{\beta} = (\tilde{\mathbf{z}}'\tilde{\mathbf{z}})^{-1}\tilde{\mathbf{z}}'\tilde{\mathbf{y}}. \tag{A.17}$$

To check for consistency, we take the probability limit of β

$$\begin{aligned} plim(\hat{\beta}) &= plim((\tilde{\mathbf{z}}'\tilde{\mathbf{z}})^{-1}\tilde{\mathbf{z}}'\tilde{\mathbf{y}}) \\ &= plim((\tilde{\mathbf{z}}'\tilde{\mathbf{z}})^{-1}\tilde{\mathbf{z}}'(\tilde{\mathbf{z}}\beta + \tilde{\mathbf{v}}^y - \beta\tilde{\mathbf{v}}^z + \mathbf{e}) \\ &= \beta + \frac{1}{\sigma^2_{\tilde{z}^*} + \sigma^2_{\tilde{\nu}^z}}(\sigma_{\tilde{\nu}^z \tilde{\nu}^y} - \beta\sigma^2_{\tilde{\nu}^z}). \end{aligned} \tag{A.18}$$

From (A.18) it can be concluded that the FE-LS estimator applied to revised data gives inconsistent results, when the parameter of interest is β in A.1. The asymptotic bias will depend not only on measurement errors in z, but also on the covariance between measurement errors in the dependent variable y and in z. Assuming that the true value for β is known, it will be possible to predict the sign and the size of this bias by computing the empirical variance $\hat{\sigma}^2_{\tilde{\nu}^z}$ and covariance $\hat{\sigma}_{\tilde{\nu}^z \tilde{\nu}^y}$ (in addition to the $\hat{\sigma}^2_{\tilde{\nu}^*}$ and $\hat{\sigma}^2_{\tilde{\nu}^z}$ terms at the denominator). Equation (A.18) encompasses the standard textbook case since, when $\sigma_{\tilde{\nu}^z \tilde{\nu}^y} = 0$, the consistency of the FE-LS estimator is only influenced by measurement errors in z, whereas measurement errors in y will affect just the estimator variance.

Albeit computationally more demanding, the analytical form of the asymptotic bias can still be derived when two or more regressors are included into the equation. Even without a formal derivation, it can be argued that relying on data contaminated by measurement errors will yield inconsistent estimates, unless the covariances between all measurement errors cancel out.

Table 1: Mean absolute value of revision errors

	(1)	(2)	(3)	(4)
$w =$	$\nu^{x}_{i,t}$	$\nu^{capb}_{1,i,t}$	$\nu^{capb}_{i,t}$	$\nu^{d}_{i,t}$
Germany	1.15	1.12	0.76	2.34
Belgium	0.68	0.64	0.46	3.40
Austria	1.13	0.79	0.97	3.85
Finland	2.51	1.25	1.30	5.38
Spain	0.66	0.58	0.77	3.83
Greece	0.73	2.94	2.83	13.59
Ireland	1.61	2.12	1.44	4.39
Italy	1.82	1.82	1.44	6.55
France	0.58	0.67	0.62	2.22
Netherlands	0.76	1.20	1.15	6.95
Portugal	1.30	1.56	1.56	4.63
Sweden	1.20	1.72	1.19	4.70
Denmark	0.75	1.18	0.96	6.19
UK	0.81	1.26	0.75	6.33
Norway	1.23	2.77	2.49	9.16
US	1.18	1.65	0.71	4.25
Canada	0.80	1.21	0.88	5.19
Japan	2.18	1.51	1.25	11.82
Australia	0.78	1.18	1.08	2.95
Mean	1.15	1.43	1.19	5.73

Source: Author own calculations based on the December Editions of the OECD Economic Outlook, from Number 56 to 80.
Note: As defined in Section (2.1), $\nu^{x}_{i,t}$,$\nu^{capb}_{i,t}$ and $\nu^{d}_{i,t}$ are the revisions errors in the current-year estimates of the $capb$,x and d respectively. $\nu^{capb}_{1,i,t}$ is the revision error in the one-year-ahead forecasts of the $capb$. Entries in the Table are the mean absolute values of revisions computed as $\sum_1^T \frac{1}{T}|w|$, where $t = 1$ corresponds to the first year of observation (1994) and T to the end of the sample (2006). Values are in percentage points.

Table 2: Empirical correlations between revision errors in the dependent variable and the ones in the regressors.

	(1)	(2)	(3)
$w =$	$\nu^{x}_{i,t-1}$	$\nu^{capb}_{i,t-1}$	$\nu^{d}_{i,t-1}$
Germany	-0.75	0.81	-0.06
Belgium	-0.21	0.52	-0.29
Austria	-0.69	0.76	-0.44
Finland	-0.70	0.30	-0.03
Spain	-0.71	0.84	0.14
Greece	-0.58	0.75	-0.86
Ireland	-0.39	0.75	-0.53
Italy	-0.91	0.95	-0.43
France	-0.55	0.74	0.07
Netherlands	-0.68	0.81	-0.15
Portugal	-0.65	0.45	-0.42
Sweden	-0.64	0.78	-0.34
Denmark	-0.67	0.77	-0.43
UK	-0.69	0.85	0.32
Norway	-0.30	0.75	0.07
US	-0.52	0.92	0.69
Canada	-0.44	0.66	0.72
Japan	-0.83	0.86	0.22
Australia	-0.36	0.62	-0.02
Mean	-0.59	0.73	-0.09

Source: Author own calculations based on the December Editions of the OECD Economic Outlook, from Number 56 (1994) to 80 (2006).
Note: Entries in the Table are $\hat{\rho}(\nu^{capb}_{1,i,t}, w)$: the empirical correlations between revision errors in the one-year-ahead forecast for the $capb$ and the revision errors in the current-year estimates (for year $t-1$) of the $capb$,x and d, respectively named $\nu^{x}_{i,t-1}$, $\nu^{capb}_{i,t-1}$ and $\nu^{d}_{i,t-1}$ as in Section (2).

Table 3: Simple fiscal rule estimates:
effects of introducing real-time information.

	Dependent variable		
	$capb_{i,t}$	$capb_{i,t}$	$capb_{i,t\|t-1}$
	(1)	(2)	(3)
$x_{i,t-1}$	-0.14**		
	-2.32		
$x_{i,t-1\|t-1}$		0.06	0.33***
		0.69	*3.60*
R^2 (within)	0.025	0.002	0.059
Observations	228	228	228
Countries	19	19	19

Notes: t statistic in italics. Estimation method: Fixed effects least squares. * significant at 10%; ** significant at 5%; *** significant at 1%.

Table 4: Estimates of the baseline fiscal policy rules: effects of introducing real-time information.

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Revised data:	$capb, x, d$		$capb, x$		$capb$		$-$	
Real-time data:	$-$		x		x, d		$capb, x, d$	
Dependent variable	$capb_{i,t}$		$capb_{i,t}$		$capb_{i,t}$		$capb_{i,t\|t-1}$	
Lagged dep.var.	$capb_{i,t-1}$		$capb_{i,t-1}$		$capb_{i,t-1}$		$capb_{i,t-1\|t-1}$	
ρ	0.69***	0.69***	0.69***	0.69***	0.68***	0.68***	0.78***	0.78***
	16.50	*16.19*	*15.76*	*15.63*	*15.60*	*15.49*	*29.37*	*29.96*
Output gap	$x_{i,t-1}$		$x_{i,t-1\|t-1}$		$x_{i,t-1\|t-1}$		$x_{i,t-1\|t-1}$	
β	-0.13***		-0.04		-0.03		0.08*	
	-2.84		*-0.63*		*-0.45*		*1.76*	
Negative output gap		$x_{i,t-1} \leq 0$		$x_{i,t-1\|t-1} \leq 0$		$x_{i,t-1\|t-1} \leq 0$		$x_{i,t-1\|t-1} \leq 0$
β_1		-0.08		-0.02		-0.01		-0.02
		-1.25		*-0.26*		*-0.11*		*-0.35*
Positive output gap		$x_{i,t-1} > 0$		$x_{i,t-1\|t-1} > 0$		$x_{i,t-1\|t-1} > 0$		$x_{i,t-1\|t-1} > 0$
β_2		-0.25**		-0.07		-0.08		0.29***
		-2.33		*-0.64*		*-0.56*		*3.38*
Debt	$d_{i,t-1}$		$d_{i,t-1}$		$d_{i,t-1\|t-1}$		$d_{i,t-1\|t-1}$	
θ	0.01	0.01	0.02***	0.02***	0.02***	0.02***	0.01	0.01
	1.59	*1.56*	*2.69*	*2.62*	*2.70*	*2.64*	*0.22*	*0.54*
Dummy EMU (ψ)	-0.23	-0.24	-0.47**	-0.47**	-0.41*	-0.42*	-0.35**	-0.34**
	-1.00	*-1.04*	*-2.03*	*-2.04*	*-1.77*	*-1.78*	*-2.29*	*-2.19*
R^2								
Within	0.59	0.59	0.57	0.57	0.57	0.57	0.83	0.83
Between	0.95	0.95	0.90	0.90	0.91	0.91	0.97	0.97
Overall	0.83	0.83	0.79	0.79	0.80	0.80	0.91	0.91
Observations	228	228	228	228	228	228	228	228
Countries	19	19	19	19	19	19	19	19

Source: author own calculations based on the December Issues of the OECD Economic Outlook from No. 56 to No. 80.

Notes: the estimated fiscal rules are equations (2), (3), (4) and (5), where positive and negative output gaps indicators are omitted for simplicity in the baseline specifications. The notation $x \leq 0$ ($x > 0$) refer to the regressor constructed as $I(x \leq 0)x$ ($I(x > 0)x$) where $I(\cdot)$ is an indicator function taking value 1 when the output gap is negative (positive) and 0 otherwise. t statistics in italics. * significant at 10%; ** significant at 5%; *** significant at 1%. Sample: 1994-2006. Estimation method: Fixed Effects Least Squares.

Table 5: Test for the number of thresholds in fiscal policy rules.
Threshold variable: output gap

	Real-time data		Revised data	
	$\hat{\rho}_1 = 0.78$	$\hat{\rho}_1 = 1$	$\hat{\rho}_1 = 0.69$	$\hat{\rho}_1 = 1$
	(1)	(2)	(3)	(4)
Test for single threshold				
F_1	10.17	9.51	2.80	0.57
p-value	0.036	0.049	0.545	0.943
critical values:				
10%	7.65	7.28	9.90	10.53
5%	9.20	9.43	12.31	13.75
1%	13.92	13.64	19.42	20.35
Test for double threshold				
F_2	4.47	3.32		
p-value	0.172	0.289		
critical values:				
10%	5.70	6.48		
5%	7.19	7.94		
1%	10.81	13.16		

Notes: $\hat{\rho}_1$ is the autoregressive coefficient from the first-stage regression (Column 1 and 3), or imposed equal to one (Columns 2 and 4). 1000 bootstrap replications are used to simulate the asymptotic distribution of the F_1 and F_2 likelihood ratio statistics.

Table 6: Single threshold fiscal rule estimates.
Threshold variable: output gap

	Real-time data	
	$\hat{\rho}_1 = 0.78$	$\hat{\rho}_1 = 1$
Threshold estimates		
$\hat{\gamma}_x$	-1.2	-1.2
68% confidence interval	[-1.5, 1.0]	[-1.5, 1.1]
95% confidence interval	[-2.4, 1.1]	[-2.4, 1.1]
Cyclical sensitivities		
$\hat{\beta}_1$	0.02	-0.07
	0.47	*-1.13*
$\hat{\beta}_2$	0.27***	0.21**
	3.63	*2.40*
Observations in regime 1	96	96
Observations in regime 2	132	132

Notes: t statistics are in italics. * significant at 10%; ** significant at 5%; *** significant at 1%.

Table 7: Test for the number of thresholds in fiscal policy rules.
Threshold variable: general government public debt as % of GDP

	Real-time data		Revised data	
	$\hat{\rho}_1 = 0.78$	$\hat{\rho}_1 = 1$	$\hat{\rho}_1 = 0.69$	$\hat{\rho}_1 = 1$
	(1)	(2)	(3)	(4)
Debt definition: gross financial liabilities. Sample: 19 OECD countries				
Test for single threshold				
F_1	7.81	4.28	5.63	4.70
p-value	0.337	0.775	0.670	0.657
critical values:				
10%	12.89	12.33	13.72	10.15
5%	15.25	14.61	15.99	12.22
1%	20.08	22.00	21.11	16.06
Debt definition: Maastricht. Sample: 11 European Union countries				
Test for single threshold				
F_1	10.67	8.16	3.96	2.50
p-value	0.122	0.350	0.807	0.948
critical values:				
10%	11.01	12.78	11.84	10.40
5%	12.82	15.31	13.84	12.28
1%	16.16	19.38	18.63	17.15

Notes 1. The 11 European Union countries included in the bottom test are Germany, Belgium, Spain, Greece, Ireland, Italy, France, Netherlands, Portugal, Denmark and the United Kingdom.
2. as in Table 5.

Table 8: Estimates of the baseline (fully-real-time) fiscal policy rule: robustness checks.

Fiscal rule specification
Columns 1,3,5,7,9: $capb_{i,t|t-1} = \alpha_i + \rho capb_{i,t-1|t-1} + \beta w + \theta d_{i,t-1|t-1} + \psi\, emu_{i,t} + \varepsilon_{i,t}$
Columns 2,4,6,8,10: $capb_{i,t|t-1} = \alpha_i + \rho capb_{i,t-1|t-1} + \beta_1 I(w \leq 0)w + \beta_2 I(w > 0)w + \theta d_{i,t-1|t-1} + \psi\, emu_{i,t} + \varepsilon_{i,t}$

	Benchmark		HP Filter		Real-time GDP		Forward looking		Election years	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)
ρ	0.78***	0.78***	0.73***	0.73***	0.78***	0.78***	0.77***	0.77***	0.78***	0.78***
	29.37	*29.96*	*26.19*	*26.12*	*29.61*	*29.37*	*27.97*	*28.11*	*29.61*	*30.23*
$w =$	$x_{i,t-1\|t-1}$		$x^{hp}_{i,t-1\|t-1}$		$gdp_{i,t-1\|t-1}$		$x_{i,t\|t-1}$		$x_{i,t-1\|t-1}$	
β	0.08*		0.10***		0.10**		0.15**		0.08*	
	1.76		*4.44*		*2.29*		*2.22*		*1.71*	
β_1		-0.02		0.08		0.03		-0.05		-0.03
		-0.35		*1.42*		*0.38*		*-0.74*		*-0.42*
β_2		0.29***		0.11***		0.16*		0.46***		0.29***
		3.38		*3.36*		*1.85*		*3.53*		*3.39*
θ	0.01	0.01	-0.01	-0.01	-0.01	-0.01	0.01	0.01	0.01	0.01
	0.22	*0.54*	*-1.53*	*-1.44*	*-0.32*	*-0.32*	*0.20*	*0.28*	*0.17*	*0.49*
emu (ψ)	-0.35**	-0.34**	-0.43***	-0.43**	-0.26*	-0.28*	-0.38**	-0.35**	-0.35**	-0.34**
	-2.29	*-2.19*	*-2.96*	*-2.92*	*-1.77*	*-1.90*	*-2.40*	*-2.27*	*-2.29*	*-2.20*
$elect$									-0.21*	-0.22**
									-1.89	*-1.96*
Method	FE-LS	FE-LS	FE-LS	FE-LS	FE-LS	FE-LS	IV	IV	FE-LS	FE-LS
R^2										
Within	0.83	0.83	0.83	0.83	0.83	0.83	-	-	0.83	0.83
Between	0.97	0.97	0.94	0.94	0.97	0.97	-	-	0.97	0.97
Overall	0.91	0.91	0.88	0.88	0.91	0.91	0.81	0.82	0.91	0.91
Sargan-p	-	-	-	-	-	-	0.73	0.88	-	-
Obs	228	228	228	228	228	228	228	228	228	228
Countries	19	19	19	19	19	19	19	19	19	19

Source: author own calculations based on the December Issues of the OECD Economic Outlook from No. 56 to No. 80.
Notes: Columns 1 and 2 report the benchmark estimates from Table 4 (fully-real-time specification). The estimates reported in Columns 3 and 4 are from regressions in which the output gap has been computed as deviation from a trend estimated by the Hodrick-Prescott filter on real GDP series in levels, measured in real-time. Columns 5 and 6 present the results when real-time values of the real GDP growth rate are used. Columns 7 and 8 present the results when the one-year-ahead forecast of the output gap is used. IV methods are employed in the estimation. The instruments used are the current-year (for year $t-1$) GDP growth and the current-year (for year $t-1$) unweighted average of the output gap over the $j \neq i$ 18 OECD countries. t statistics are in italics. * significant at 10%; ** significant at 5%; *** significant at 1%. Sample: 1994-2006.

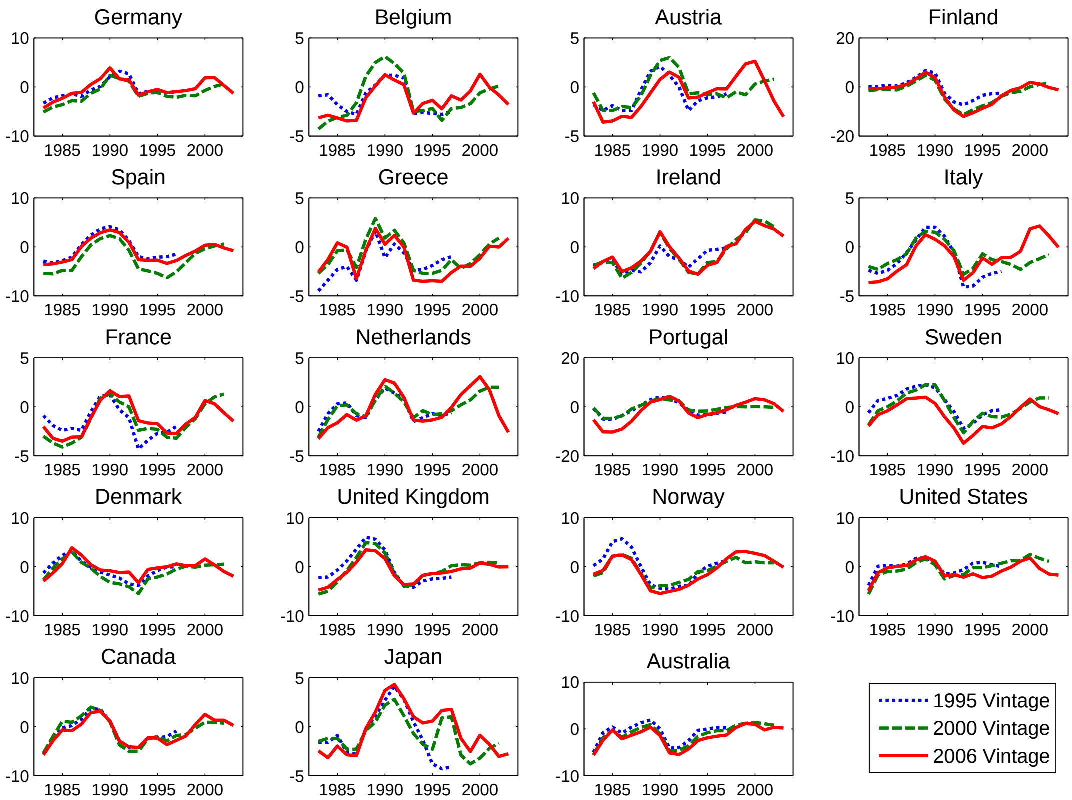

Figure 1: Output gaps for some OECD countries from three data vintages. Data sources: OECD Economic Outlook No. 58 (December 1995), No. 68 (December 2000) and No. 80 (December 2006).

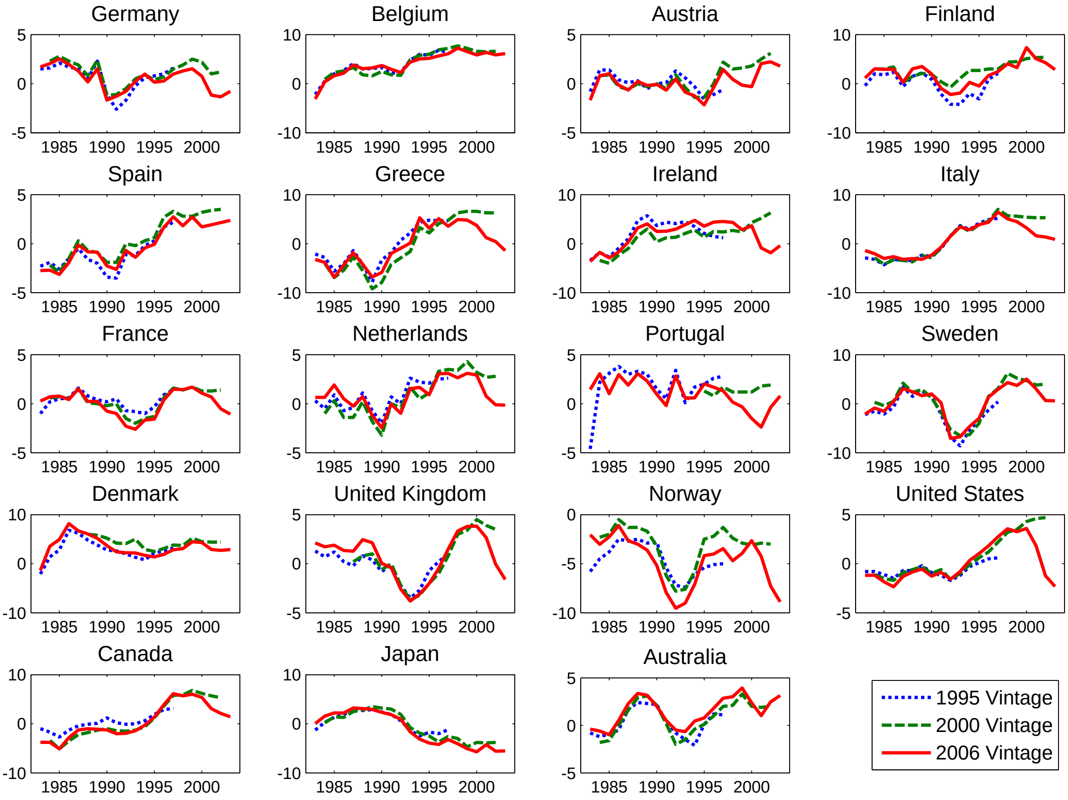

Figure 2: Cyclically-adjusted primary balances as percentage of potential GDP for some OECD countries from three data vintages. Data sources: OECD Economic Outlook No. 58 (December 1995), No. 68 (December 2000) and No. 80 (December 2006).

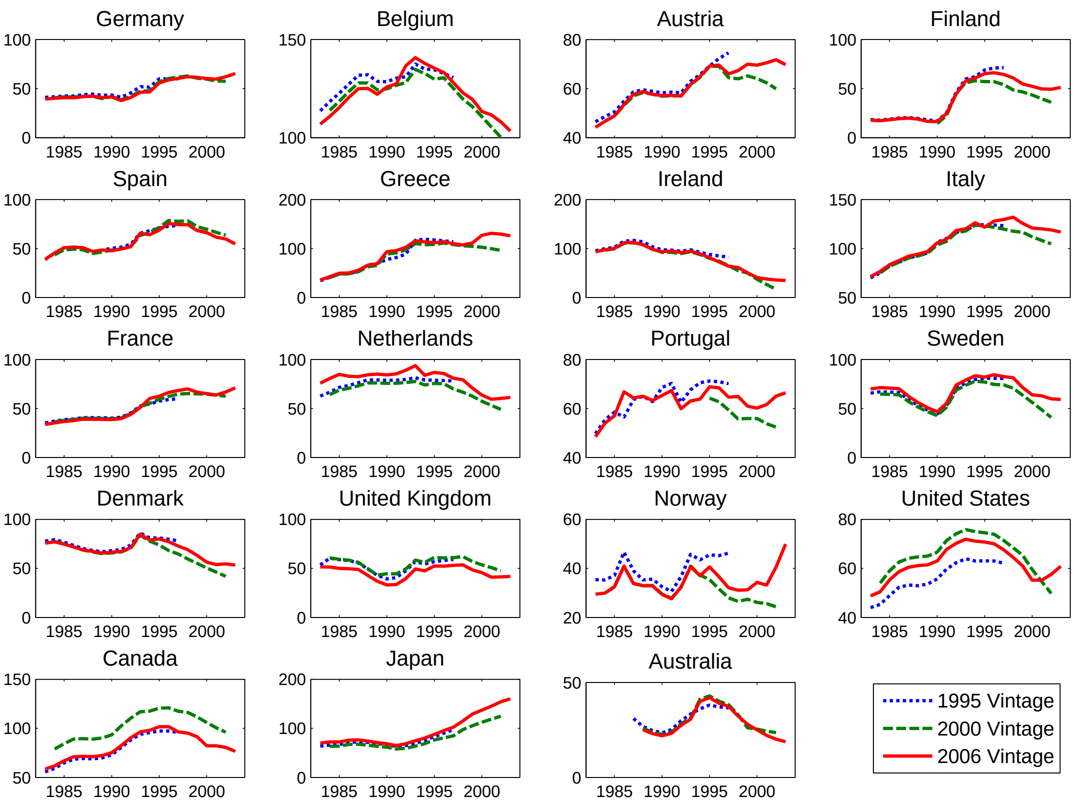

Figure 3: General government gross financial liabilities as percentage of GDP for some OECD countries from three data vintages. Data sources: OECD Economic Outlook No. 58 (December 1995), No. 68 (December 2000) and No. 80 (December 2006).

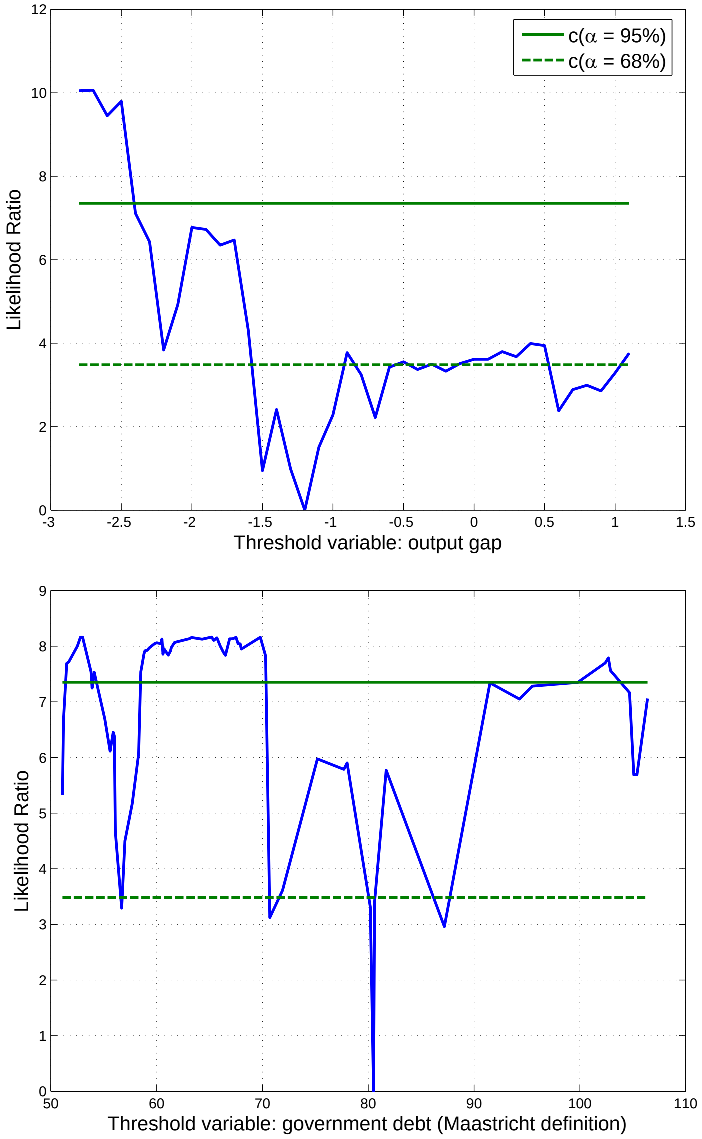

Figure 4: Confidence interval construction in single threshold models (equation (23) and (24)). Blue lines: likelihood ratio statistic LR_1 (equation (12)); green solid lines: 95% critical value (10.59); green dotted line: 68% critical value (7.35). The graphs refer to the case $\hat{\rho}_1 = 0.78$ (top panel) and $\hat{\rho}_1 = 1$ (bottom panel), when real-time data are used in estimation. Sample: 1994-2006. Top graph: 19 OECD countries, bottom graph: 11 European Union countries.

List of working papers released by CEPII[1]

No	*Title*	*Authors*
2007-09	Global Ageing and Macroeconomic Consequences of Demographic Uncertainty in a Multi-regional Model	J. Alho & V. Borgy
2007-08	The Effect of Domestic Regulation on Services Trade Revisited	C. Schwellnus
2007-07	The location of domestic and foreign production affiliates by French multinational firms	T.Mayer I. Méjean B. Néfussi
2007-06	Specialisation across Varieties within Products and North-South Competition	L. Fontagné, G. Gaulier & S. Zignago
2007-05	Trade Costs and the Home Market Effect	M. Crozet & F. Trionfetti
2007-04	The Impact of Regulations on Agricultural Trade: Evidence from SPS and TBT Agreements	A.-C. Disdier, L. Fontagné & M. Mimouni
2007-03	International Comparisons of Living Standards by Equivalent Incomes	M. Fleurbaey & G. Gaulier
2007-02	Does Risk Aversion Drive Financial Crises? Testing the Predictive Power of Empirical Indicators	V. Coudert & M. Gex
2007-01	Asian Catch Up, World Growth and International Capital Flows in the XXIst Century : A Prospective Analysis with the INGENUE 2 Model	M. Aglietta, V. Borgy, J. Château, M. Juillard, J. Le Cacheux, G. Le Garrec & V. Touzé
2006-27	Current Account Reversals and Long Term Imbalances: Application to the Central and Eastern European Countries	K. Benhima & O. Havrylchyk
2006-26	On Legal Origins and Brankruptcy Laws: the	J. Sgard

[1] Working papers are circulated free of charge as far as stocks are available; thank you to send your request to CEPII, Sylvie Hurion, 9, rue Georges-Pitard, 75015 Paris, or by fax : (33) 01 53 68 55 04 or by e-mail Hurion@cepii.fr. Also available on: \\www.cepii.fr. Working papers with * are out of print. They can nevertheless be consulted and downloaded from this website.

[1] Les documents de travail sont diffusés gratuitement sur demande dans la mesure des stocks disponibles. Merci d'adresser votre demande au CEPII, Sylvie Hurion, 9, rue Georges-Pitard, 75015 Paris, ou par fax : (33) 01 53 68 55 04 ou par e-mail Hurion@cepii.fr. Egalement disponibles sur : \\www.cepii.fr. Les documents de travail comportant * sont épuisés. Ils sont toutefois consultable sur le web CEPII.

	European Experience (1808-1914)	
2006-25	Taux d'intérêt et marchés boursiers : une analyse empirique de l'intégration financière internationale	V. Borgy & V. Mignon
2006-24	Changing Patterns of Domestic and Cross-Border Fiscal Policy Multipliers in Europe and the US	A. Bénassy-Quéré & J. Cimadomo
2006-23	Market Access Impact on Individual Wage: Evidence from China	L. Hering & S. Poncet
2006-22	FDI in Chinese Cities: Spillovers and Impact on Growth	N. Madariaga & S. Poncet
2006-21	Taux d'intérêt et marchés boursiers : une analyse empirique de l'intégration financière internationale	V. Borgy & V. Mignon
2006-20	World Consistent Equilibrium Exchange Rates	A. Bénassy-Quéré, A. Lahrèche-Révil & V. Mignon
2006-19	Institutions and Bilateral Asset Holdings	V. Salins & A. Bénassy-Quéré
2006-18	Vertical Production Networks: Evidence from France	M. Fouquin, L. Nayman & L. Wagner
2006-17	Import Prices, Variety and the Extensive Margin of Trade	G. Gaulier & I. Méjean
2006-16	The Long Term Growth Prospects of the World Economy: Horizon 2050	S. Poncet
2006-15	Economic Integration in Asia: Bilateral Free Trade Agreements Versus Asian Single Market	M. H. Bchir & M. Fouquin
2006-14	Foreign Direct Investment in China: Reward or Remedy?	O. Havrylchyk & S. Poncet
2006-13	Short-Term Fiscal Spillovers in a Monetary Union	A. Bénassy-Quéré
2006-12	Can Firms' Location Decisions Counteract the Balassa-Samuelson Effect?	I. Méjean
2006-11	Who's Afraid of Tax Competition? Harmless Tax Competition from the New European Member States	A. Lahrèche-Révil
2006-10	A Quantitative Assessment of the Outcome of the Doha Development Agenda	Y. Decreux & L. Fontagné
2006-09	Disparities in Pension Financing in Europe: Economic and Financial Consequences	J. Château & X. Chojnicki
2006-09	Disparities in Pension Financing in Europe: Economic and Financial Consequences	J. Château & X. Chojnicki
2006-08	Base de données CHELEM-BAL du CEPII	H. Boumellassa & D. Ünal-Kesenci

2006-07	Deindustrialisation and the Fear of Relocations in the Industry	H. Boulhol & L. Fontagné
2006-06	A Dynamic Perspective for the Reform of the Stability and Gowth Pact	C. Deubner
2006-05	China's Emergence and the Reorganisation of Trade Flows in Asia	G. Gaulier, F. Lemoine & D. Ünal-Kesenci
2006-04	Who Pays China's Bank Restructuring Bill?	G. Ma
2006-03	Programme de travail CEPII pour 2004	
2006-02	Technology Differences, Institutions and Economic Growth : a Conditional Conditional Convergence	H. Boulhol
2006-01	Croissance et régimes d'investissement	P. Villa

CEPII
DOCUMENTS DE TRAVAIL / WORKING PAPERS

Si vous souhaitez recevoir des Documents de travail,
merci de remplir le coupon-réponse ci-joint et de le retourner à :

Should you wish to receive copies of the CEPII's Working papers,
just fill the reply card and return it to:

Sylvie HURION – Publications
CEPII – 9, rue Georges-Pitard – 75740 Paris – Fax : (33) 1.53.68.55.04
sylvie.hurion@cepii.fr

M./Mme / Mr./Mrs ....

Nom-Prénom / Name-First name ....

Titre / Title ....

Service / Department...

Organisme / Organisation ....

Adresse / Address..

Ville & CP / City & post code..
Pays / Country .... Tél. ....
Your e-mail ....

Désire recevoir les **Document de travail** du CEPII n° :

Wish to receive the ***CEPII's Working Papers No****:* ....

....

....

....

....

....

....

....

....

☐ Souhaite être placé sur la liste de diffusion permanente (**pour les bibliothèques**)
Wish to be placed on the standing mailing list ***(for Libraries)****.*