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by Jacopo Cimadomo[2]

In 2008 all ECB publications feature a motif taken from the €10 banknote.

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Abstract

This paper argues 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 was planned by fiscal authorities in the budget law. When fiscal policy rules are estimated on real-time data, our results indicate that OECD countries often planned a counter-cyclical fiscal stance, especially during economic expansions, whereas conventional findings based on revised data point towards pro-cyclicality. This finding calls into question the effectiveness of discretionary fiscal policies to fine tune the business cycle, as (pro-cyclical) actual outcomes tend to deviate from (counter-cyclical) fiscal plans. Furthermore, we test whether threshold effects might be at play in the reaction of fiscal policy to the economic cycle and to public debt accumulation. It emerges that the intended 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.

Keywords: Fiscal policy, real-time data, revision errors, endogenous threshold models.
JEL Classification: C23, E30, E62, H30, H60

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*b*) and Kaminsky, Reinhart and Vegh (2004)). Otherwise stated, budgetary decisions such as tax increases and retrenchments of 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, and 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 variables 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", i.e. 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.

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. However, from the 1950s on, and especially during the 1970s and the 1980s, 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";[1] ii) there are long and uncertain time lags in the implementation of fiscal measures; iii) institutional constraints may prevent a timely use of fiscal policy; iv) fiscal policy decisions are, often, irreversible.[2]

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 first 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. Lane (2003*a*) shows that the "pro-cyclicality bias" is more severe for developing countries than for developed ones. As for the OECD economies, Lane (2003*c*) argues that spending categories are characterized by different levels of cyclicality and that countries with volatile output and dispersed political power are most likely affected

[1] 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.

[2] 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 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 of "close-to-balance or in surplus" (see Brunila, Buti and in't Veld (2002) and Buti and Sapir (2006)). In this context, discretionary fiscal measures are recommended only to the extent that excessive deficits need to be corrected and to reach balanced budget positions. In particular, the new version of the Pact (as from the ECOFIN Council of March 2005) lays down that Member States characterized by excessive deficits should pursue annual improvements of their cyclically-adjusted balances, net of one-off and other temporary measures, of at least 0.5 of GDP as benchmark.

by pro-cyclical fiscal policies. Focusing on the same set of countries, the OECD (2003) emphasizes that the stance of fiscal policy tends to be predominantly counter-cyclical in "bad" times and pro-cyclical in "good" ones. Similar results emerge from a study by the European Commission (2004) on euro area countries, thereby suggesting that the EU fiscal framework would have not helped to eradicate the occurrence of a pro-cyclical bias during booming economic periods.

Summing up, the empirical evidence from this literature seems to be quite consensual as regards developing countries, pointing to 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.

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 made, and in particular when it is supposed that the updated, revised information is available *ex-ante* to decision makers, the analysis of policy-makers behavior may be drastically misleading.[3]

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)). However, 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 important exception is the paper by Forni and Momigliano (2005). These authors estimate, for a panel of OECD countries, fiscal policy rules linking changes of a discretionary fiscal policy indicator, the cyclically-adjusted primary budget balance as percentage of potential GDP, 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.[4]

[3] 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.

[4] 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, the effects of automatic stabilizers and discretionary measures cannot be disentangled. More recently, Golinelli and Momigliano (2006) include real-time observations of the primary balance on the right-hand side of their regression equation, but they use revised data for dependent variable.

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 policy "instrument" itself. Typically, in fact, in each autumn of year $t-1$, fiscal authorities approve the budget for year t. Budget laws are mainly designed on the basis of *ex-ante* projections on the state of the economy, and of the perceived evolution of inherited public deficits and debts. In addition, the realization of planned fiscal measures depends importantly on implementations lags. Therefore, there might be relevant discrepancies between discretionary fiscal measures as approved *ex-ante* and what is observed several periods after decisions have been taken.[5]

In this paper, we gauge the "intentional" stance of fiscal policy in OECD countries based on an information set which closely mimic the one available to governments at the time of fiscal planning. Data sources are past issues of the OECD Economic Outlook since, as spelled out more in detail below, this has several advantages compared to the use of revised data, of "official" governments' projections, and of projections released by other international institutions. To the best of our knowledge, this is the first attempt to incorporate real-time observations on *all* the "ingredients" typically used in the estimation of fiscal policy rules.[6] In particular, fiscal plans as reported at the time of budgeting are employed.[7] Specifically, the following contributions are put forward:

Firstly, a real-time annual dataset is constructed by collecting data on some key fiscal and macroeconomic variables published in the December Issues of the OECD Economic Outlook from 1994 (Volume 56) to 2006 (Volume 80), for 19 OECD countries.

Second, simple fiscal policy rules, relating the cyclically-adjusted primary balance as indicator of discretionary fiscal policies, to the output gap, are estimated for this set of industrialized countries over the 1994-2006 period. It is shown that the inclusion of real-time observations of the *capb* revert the sign of the estimated parameter representing the cyclical sensitivity of discretionary policies: estimates based on *ex-post* data point to pro-cyclicality, as conventional, whereas estimates stemming from real-time data indicate counter-cyclicality. In addition, it is shown that the bias incurred in estimating a simple fiscal policy rule using revised data, when the intentional fiscal policy stance is the object interest, can be accurately predicted based on the empirical covariances between revision

[5]Note that this issue is not relevant for monetary policy analysis. In fact, under normal circumstances, central bankers can control 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. On the contrary, the cyclically-adjusted budget balance, as indicator of the discretionary component of fiscal policy, 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.

[6]See also Cimadomo (2007) for an earlier version of this work.

[7]The proposed approach has been recently followed by Giuliodori and Beetsma (2008) to gauge how budgetary decisions in one country may be affected by fiscal plans announced in other countries, and by Beetsma and Giuliodori (2008) to test how OECD governments have reacted to unexpected information about the business cycle.

errors in $capb's$ and in output gaps, and on other second-order moments.

Third, more encompassing fiscal policy rules are estimated, where movements in the fiscal policy indicator are assumed to depend not only on cyclical conditions, but also on debt developments and on a set of other control variables, similarly to 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. Moreover, when positive and negative output gaps are included separately it emerges that, *ex-post*, fiscal policy seems pro-cyclical in expansions and a-cyclical in recessions. However, when real-time data are used, the stance is gauged to be strongly counter-cyclical during upturns and a-cyclical in slowdowns.

Finally, we explore whether the discretionary reaction of fiscal policy to the business cycle and to debt accumulation is characterized by multiple regimes. A two-stage procedure applied to the Hansen's (1999) threshold panel regression model is proposed. Based on this approach, it is found that the hypothesis of a switch in the *ex-ante* cyclical behavior of fiscal policy (from acyclicality to counter-cyclicality) occurring when GDP is close to its equilibrium level cannot be 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 discuss the advantages of using cyclically-adjusted balances published in "real-time" by the OECD as proxies for intentional fiscal plans. Section 3 presents the data. Section 4 is devoted to illustrate the fiscal policy reaction functions used in the empirical analysis and the estimation results. In particular, Section 4.1 shows our finding for simple fiscal rules, Section 4.2 describes how revision errors in variables may affect estimation results, Section 4.3 focuses on baseline fiscal policy rules' estimates and Section 4.4 investigates the possible presence of multiple regimes in the conduct of fiscal policy. Finally, Section 5 reports some robustness exercises and Section 6 concludes.

2 How should we measure "ex-ante" fiscal plans?

Changes in governments' revenues and expenditures and, as a consequence, in budget balances, are driven by two main sources. On the one hand, business cycle fluctuations affect budget balances through the operation of automatic stabilizers.[8] Automatic stabilization implies that we would observe movements in balances even in absence of any legislative intervention by governments.

On the other hand, fiscal authorities may decide to actively intervene on public accounts

[8]Examples of automatic stabilizers are, on the revenue side, corporate and individual tax receipts that automatically vary due to changes in income, for given tax rates and a given definition of the tax base. Among government expenditure items, unemployment compensation is the most influenced by output fluctuations.

by, for instance, cutting tax rates or increasing public expenditures. The bulk of such *discretionary* interventions is generally incorporated in the budget law for year t as enacted by national parliaments in (autumn of) year $t-1$.[9]

Typically, two approaches have been followed to measure the discretionary component of budget balances, or of other fiscal policy indicators. The recent "narrative" approach, proposed by Romer and Romer (2007) for the analysis of the effects of tax reductions on the U.S. economic activity, is based on the idea of collecting single episodes of policy change and to record the timing and the magnitude of their (expected) effects, as reported by official documents (i.e. past budget laws, Economic Reports of the President, Congressional Records). This method, since based on governmental historical sources, should be effective in capturing policymakers' discretionary plans, or *intentions*. However, fiscal decisions as planned in the past may fail to (fully) materialize, due for example to differences between the output growth as perceived at the time of budgeting and the actual one.[10] Therefore, what we observe *ex-post*, in budget balances, could turn out to be considerably different from what planned in the past, and captured by the record of legislated measures. In other words, such an approach could be inappropriate to assess the "realized", as opposed to *ex-ante*, stance of fiscal authorities.

Alternatively, the discretionary component of fiscal policies may be isolated through cyclical-adjustment methods. Once the effects due to automatic stabilization (and interest payments) are netted out from headline nominal balances, the residual (or "structural") component, should capture the discretionary stance of fiscal authorities. Clearly, if the *intentional* stance of fiscal authorities is under investigation, cyclical-adjustment procedures should be applied to budget balances as reported when decisions have been taken, using *ex-ante* projections of output gaps. In this case, the resulting cyclically-adjusted component of the balance should closely track the record of discretionary measures as planned in the past.

Hence, in this paper we propose to use the cyclically-adjusted (primary) balance for years' t as reported in years' $t-1$ December editions of the OECD Economic Outlook as *proxy* of the intentional, or *ex-ante*, stance of fiscal policy authorities in 19 OECD coun-

[9]To be noted, discretionary measures may be decided for a variety of reasons. For example, governments may want to dampen output fluctuations, in case they deem that automatic stabilizers do not work efficiently enough. Moreover, they might react to inherited budget deficits and debts, especially in those countries affected by a system of fiscal rules, and to the extent that they are concerned about the long-term sustainability of public finances. In addition, governments may intervene on the "quality" of public expenditure and taxation, for example by allocating resources on items more likely to promote long-term economic growth and job creation. An upcoming election, and further factors, may also induce government to implement discretionary measures.

[10]For example, the estimated impact, in terms of revenue effects, of a revision of income tax brackets or marginal tax rates depends crucially on the actual evolution of the tax base. This, as depends on the state of the economy, might be substantially different from government's projections. Moreover, since governments often design their discretionary interventions in terms of *ratios* to GDP, outcomes might be sensibly different from plans as long as GDP projections result to be inaccurate.

tries. This approach has several advantages relative to alternative methods. In particular, implementing Romer and Romer's (2007) approach would be a daunting task in the present context, since many countries are involved in the analysis and a history of legislated policy changes should be gathered for each of them. Moreover, employing OECD data should improve cross-country comparability, since variables' definitions and accounting standards are homogeneous. As an alternative, one could use cyclically-adjusted balances as published by official authorities, as for example the ones reported in Stability and Convergence Programmes. However, as documented by several authors (see e.g. Annett (2006), Jonung and Larch (2004), Strauch, Hallerberg and von Hagen (2004)), these data are often affected by a "political bias" inducing an overly optimistic outlook on the state of public finances. In this respect, projections provided by the OECD are likely to be more reliable since, although the original sources of OECD data are still national statistical offices, the OECD staff applies judgment and filters out unrealistic figures when compiling its statistics. Indeed, OECD fiscal forecasts are generally found to be unbiased, as systematic under or over-predictions of budgetary figures are rare (see in particular Artis and Marcellino (2001)). An additional advantage of using OECD data is related to the timing of their release. In fact, OECD projections are published in December and, as a consequence, they are likely to reflect the full set of discretionary measures incorporated in national budget laws, which are generally approved at the end of the year by national parliaments. Projections from other institutional institutions, such as the IMF or the European Commission, could also be employed. Yet, we opt for the OECD as reference source of data for two main reasons. First, although the methodology employed to net out the effects of the economic cycle from fiscal aggregates, and the elasticities used to gauge cyclical sensitivity of single budget items, may change across institutions, cyclically-adjusted indicators published by the IMF and the European Commission are broadly in line with the ones released by the OECD. Second, the availability of several past issues of the OECD Economic Outlook makes the construction of a relatively large real-time dataset feasible.

In the light of these arguments, real-time OECD data on *capb*'s, but also on other key fiscal and macroeconomic variables, are gathered. This should enable to plausibly mimic the information set available to policymaker at the time budgetary decisions have been taken.

3 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 the previous sixteen years, "estimates" for the current-year and "forecasts" for years $t+1$ and $t+2$. Our 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. Data for three reference indicators have been collected: the output gap (deviation of actual GDP from potential GDP as percentage of potential GDP), defined hereafter as x[11], the general government debt (gross financial liabilities) to nominal GDP ratio, defined as d, and the cyclically-adjusted primary balance as percentage of potential GDP, defined as *capb*. Our real-time dataset is built on by inputting the current-year estimates, and the one year-ahead forecasts, of these key variables. Furthermore, current-year estimates of real GDP annual growth rates and of general government financial balances as percentage of nominal GDP have also been gathered. These additional indicators will be used as control variables in the baseline estimations (financial balances) and in the proposed robustness exercises (real GDP growth rates).[12]

Given a generic variable z, and a country i, we follow this notation: $z_{i,t|t-1}$ denotes the one-year-ahead forecast of z, for country i, as estimated in vintage $t-1$; $z_{i,t-1|t-1}$ indicates the current-year estimate of z, released in vintage $t-1$ and $z_{i,t}$ refers to revised data which, in the present context, are the ones from the December 2006 vintage. To be noted, the discretionary fiscal policy "instrument" used in the paper is $capb_{i,t|t-1}$, i.e. the one-year ahead forecast of the *capb*.

The relationships between real-time and revised observations of some key variables used in the rest of the paper are defined as follows

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

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*. Notice 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.[13]

[11] 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. In addition, 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 data on net debt interest payments to the ones on structural balances.

[12] Overall, more than 25.000 observations have been recorded from electronic or hard copies of the OECD Economic Outlook.

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

Figure 1, 2 and 3 display, as illustrative examples, two different vintages of data (2000 and 2006) for the variables of main interest. Even from a simple visual inspection, it can be noted that the data from first vintage 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 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 revisions are generally large. For instance, these amount to 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 forecasts and the current-year estimates of the *capb* (Columns 2 and 3, Table 1). The fourth Column of the Table, and Figure 3, indicate that the level of debt-output ratio as 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 statistical definitions and accounting rules.[14] 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 one-year ahead forecasts of the *capb*, and revision errors in current-year estimates of the output gap, the *capb* and the debt to GDP ratio. 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 $\nu_{1,i,t}^{capb}$ and $\nu_{i,t-1}^{d}$ are less uniform across countries, and approximately distributed around a zero mean value.

[14]For instance, in the case of Canada, a vertical shift of around 20 percentage points of GDP is observed between the 2000 vintage and the 2006 one, as shown in Figure 3. This is explained by a change in the statistical definition of general government gross financial liabilities for Canada, occurring in 2002. In fact, before 2002, funded government employees pension liabilities were included in this indicator whereas starting from 2002 these have been netted out by the OECD, for consistency with other countries.

4 Assessing the stance of fiscal policy in real-time

4.1 Simple fiscal policy rules

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, as represented by the *capb*, to deviations of actual output from its equilibrium level.

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*, 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.

To evaluate the potential impact of making "unreasonable" assumptions regarding the timeliness of information available to the fiscal policymaker, when the interest is on the *ex-ante* stance of governments, we proceed as follows: as a starting step, we estimate a simple fiscal policy rule relating the *capb* to the (lagged) output gap. Lagged values of the output gap are used to avoid possible endogeneity problems stemming from the fact that fiscal policy measures may affect the real activity in the same period. The sample includes the whole panel of 19 OECD countries, and ranges over the 1994-2006 period. First, only revised data (i.e. the ones from the December 2006 vintage) are used in the estimation:

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

Then, current-year (real-time) estimates of the output gap are employed:

$$capb_{i,t} = \alpha_i + \beta x_{i,t-1|t-1} + \varepsilon_{i,t}. \tag{2}$$

Finally, real-time figures for both the *capb* and x are incorporated:

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

In the equations above, (unobserved) country-specific effects are captured by the terms α_i, which should also help to control for systematic biases leading to over (under) reporting of fiscal stances. Table 3 shows the results, based on Fixed-Effect (Within) Least-Squares (FE-LS). The estimated cyclical sensitivity parameter shown 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 of 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. Interestingly, the estimated regression slope turns positive, indicating counter-cyclicality, and it is significant at the 1% level.[15] This implies that assessments on the *intentional* stance of fiscal authorities based on the estimation of simple fiscal rules could be highly misleading when updated information, in the form of revised data, is used.

4.2 Bias prediction

The discretionary fiscal policy indicator, i.e. the *capb*, and the explanatory variable, i.e. the output gap, incorporated in a simple fiscal policy rule as the one presented above, are contaminated by large measurement errors (see Table 1), which seem to be also highly cross-correlated (see Table 2). 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 (LS) 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 Instrumental Variable methods, provided that appropriate instruments are found. Under standard assumptions, on the other hand, the presence of measurement errors in the *dependent* variable impacts on the variance of LS estimators, but not on their consistency. This holds also in panel regressions. Therefore, applying the FE-LS estimator to model (3) above, where the one-year-ahead forecast of the *capb* is included in the left-hand-side of the regression equation, should asymptotically (for $N, T \rightarrow \infty$) yield the same results (in terms of point estimates) as when revised data are used for the dependent variable.

Here it is shown that when the conventional assumption of uncorrelatedness between measurement errors is dropped, and this seems to be in line with the results presented in Table 2, the FE-LS estimator becomes biased and inconsistent not only due to measurement errors in the regressor, as in the standard case, but also to the fact that revision errors in the dependent variable and the independent one 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

[15] 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 of 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.

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 addressed yet in the literature on panel data.

Let us consider a simple fiscal policy rule as in Section 4.1, 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

$$capb_{i,t|t-1} = \alpha_i + \beta x_{i,t-1|t-1} + \varepsilon_{i,t}, \tag{4}$$

where $\varepsilon_{i,t} \sim i.i.d(0, \sigma^2_{\varepsilon})$. Suppose that the "true", or real-time, values of the *capb* and x are not observed. Instead, we observe the revised data, denoted by $capb_{i,t}$ and $x_{i,t-1}$. 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 $\nu^{capb}_{1,i,t}$ and $\nu^{x}_{i,t-1}$ to be contemporaneously correlated, Appendix A formally shows that the asymptotic bias incurred in estimating (4) by FE-LS and using revised information is equal to

$$BIAS = \frac{1}{\sigma^2_{\tilde{x}^*} + \sigma^2_{\tilde{\nu}^x}}(\sigma_{\tilde{\nu}^x\tilde{\nu}_1^{capb}} - \beta\sigma^2_{\tilde{\nu}^x}), \tag{5}$$

where $\sigma^2_{\tilde{x}^*}$ is the variance of the "true" values of x, pooled across groups after removing individual means; $\sigma^2_{\tilde{\nu}^x}$ 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.[16] Equation (5) 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.

Replacing the empirical counterparts of the theoretical second-order moments included in (5), and assuming that the "true" β is the one obtained from the real-time regression, it will be possible to accurately predict the size and the direction of the bias. It particular, the estimated bias will be equal to

$$BI\hat{A}S = \frac{1}{\hat{\sigma}^2_{\tilde{x}^*} + \hat{\sigma}^2_{\tilde{\nu}^x}}(\hat{\sigma}_{\tilde{\nu}^x\tilde{\nu}_1^{capb}} - \beta\hat{\sigma}^2_{\tilde{\nu}^x}) = -0.53,$$

where $\hat{\sigma}^2_{\tilde{x}^*} = 1.54$, $\hat{\sigma}^2_{\tilde{\nu}^x} = 1.75$, $\hat{\sigma}_{\tilde{\nu}^x\tilde{\nu}_1^{capb}} = -1.14$ and $\beta = 0.35$. To be noted, by adding the estimated (negative) bias to the "true" β we get -0.18, a value very close to the regression slope obtained from revised data and equal to -0.14. This suggests that

[16] Note that the derivation of the asymptotic bias reported in 5 is robust to the assumption that revision errors might be autocorrelated, as shown in Appendix A.

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) implying a pro-cyclical stance.[17] As a byproduct, the relative contribution to the overall bias of the revision errors in x, and of 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.

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. Yet, 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.

4.3 Baseline fiscal policy rules

Simple fiscal policy rules as the ones presented in Section 4.1 may suffer from an omitted variable problem, since fiscal authorities are likely to react not only to cyclical developments, but also to current deficits, to the debt to GDP ratio as indicator of long-term fiscal sustainability (see in particular Bohn (1998)), and to other possible indicators. Therefore, the latest generation of fiscal policy rules includes a (relatively) large set of explanatory variables to account 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.

In the light of these arguments, we assess the impact of incorporating real-time information based on a more encompassing fiscal policy reaction function, similar to the one proposed by Galì and Perotti (2003). As before, a "backward-looking" specification of the rule is proposed, to avoid possible feedbacks stemming from the fact that fiscal measures may impact on output contemporaneously.[19] A battery of four fiscal policy rules is estimated, where the amount of real-time information is progressively increased, from a "fully-revised" scenario to a "fully-real-time" one:

[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] As a robustness exercise, fiscal policy rules including contemporaneous output gaps (or one-year ahead forecasts) as independent variables are also estimated, as shown in Section 5.

i) Revised-data: $capb, x, d, emu$; no real-time data ("fully-revised" rule);

$$capb_{i,t} = \alpha_i + \rho capb_{i,t-1} + \beta x_{i,t-1} + \theta d_{i,t-1} + \psi\, emu_{i,t-1} + \varepsilon_{i,t}. \tag{6}$$

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

$$capb_{i,t} = \alpha_i + \rho capb_{i,t-1} + \beta x_{i,t-1|t-1} + \theta d_{i,t-1} + \psi\, emu_{i,t-1} + \varepsilon_{i,t}. \tag{7}$$

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

$$capb_{i,t} = \alpha_i + \rho capb_{i,t-1} + \beta x_{i,t-1|t-1} + \theta d_{i,t-1|t-1} + \psi\, emu_{i,t-1|t-1} + \varepsilon_{i,t}. \tag{8}$$

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

$$capb_{i,t|t-1} = \alpha_i + \rho capb_{i,t-1|t-1} + \beta x_{i,t-1|t-1} + \theta d_{i,t-1|t-1} + \psi\, emu_{i,t-1|t-1} + \varepsilon_{i,t}. \tag{9}$$

where emu is a dummy variable which equals one in years when the total deficit exceeded the 3% limit imposed by the Maastricht Treaty and the Stability and Growth Pact, for the countries having joined the European Monetary Union, and zero otherwise. The idea is that, during the convergence process (i.e. before 1997), EMU governments may have had incentives to curb public deficits when they were higher than the 3% threshold, to achieve the goal of the admission to the Monetary Union. From 1997 on, they might have wanted to avoid the sanctions implied by the Pact. These effects would be captured by an estimated positive (and significant) ψ. In models (6) and (7) emu has been computed by using revised data of deficit-GDP ratios whereas real-time data of this indicator are used in models (8) and (9).[20]

For each of the models presented above, two additional variables are constructed by interacting x with a dummy indicator 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, these regressors (defined as "negative" and "positive" output gaps) are included separately in the proposed fiscal rules, where the associated coefficients are β_1 and β_2.

The panel specification underlying the proposed regression models implies that policy-makers in OECD countries behave uniformly, as far as reactions to output fluctuations and

[20]Models (8) and (9) 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.

debt dynamics are concerned. Therefore, β and θ gauge, respectively, the common-across-countries "cyclical sensitivity" and "sustainability concern" of fiscal authorities.[21] Such a homogeneity assumption is made necessary to avoid degree-of-freedom-related problems in estimation, although it cannot be discarded that fiscal authorities may follow (at least partially) different reaction functions. However, a fixed-effects panel model approach has the advantage that possible (unobserved) heterogeneities in the conduct of fiscal policy are captured by the coefficients α_i.

Table 4 reports FE-LS estimates of model (6) through (9).[22] Column 1 and 2 display the results when only ex-post data are used in the regressions. The cyclical sensitivity parameter estimate is -0.15 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 $\hat{\beta}_1$ 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 and on deficit-GDP ratios used to computed the *emu* dummy variable does not alter the picture much, except that the $\hat{\psi}$ turns significant (Columns 5 and 6). This suggests that, in the years preceding and following the adoption of the single currency, EMU countries have attempted to reduce deficits (the *capb* rises) when these have exceeded the 3% ceiling.

When the one-year-ahead forecasts of the *capb* and the current-year estimates of it are used in the regression, 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. The "unconditional" cyclical sensitivity of the budget authority (i.e. β) can be thought as a (weighted) average of the cyclical sensitivities, conditional on positive and negative phases of the business cycle. Then, as results in Column 8 indicate, a very strong counter-cyclical reaction during expansions plays a key role in explaining the positive $\hat{\beta}$, whereas fiscal policy seems to remain acyclical in recessions. In particular, during upturns, a one percentage point increase of the output gap induces a 0.24 percentage point

[21] 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.

[22] 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.

(planned) fiscal tightening. 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 generally found in papers using revised data (see e.g. Bohn (1998) and European Commission (2004)), since $\hat{\theta}$ is not statistically different from zero. It is also worth noticing that persistence in fiscal planning plays a key role in accounting for $capb's$ dynamics, as documented by $\hat{\rho}$ being highly significant and close to one. Finally, the R^2 of the "fully-real-time" regression is remarkably higher compared to ones of models (6) through (8), thereby indicating the appropriateness of this specification when the interest is on the *intentional* fiscal policy stance.[23]

4.4 Multiple regimes in fiscal policy

The issue of whether fiscal policy behaves asymmetrically along the business cycle has been addressed in the previous Section by relying on the notion of "good times" as periods of positive output gaps and "bad" ones as years in which the actual output was below the potential one. The same approach is followed by many other papers on fiscal policy rules (see e.g. Gavin and Perotti (1997), European Commission (2004) and OECD (2003)).

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, implying 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 (different from zero). Moreover, the conduct of fiscal policy may be characterized by more than two regimes. In particular, Manasse (2005) suggests that the cyclical sensitivity of fiscal policy may vary when slowdowns are particularly severe and upturns strong, compared to intermediate states of the business cycle.[24]

[23] On the cyclical response of fiscal policy, Galì and Perotti (2003), using the (then-available) *ex-post* data from the December 2002 OECD Economic Outlook, argue that the stance of fiscal policy has been a-cyclical in EMU countries, and counter-cyclical in OECD non-EMU ones, over the 1992-2001 period. The vintage of data used in that work, and in particular its proximity to the 1999-2000 expansion, may help to reconcile Galì and Perotti's (2003) findings with the ones of this paper. In fact, the 1999-2000 upturn was characterized by large *downward* revision of potential outputs, and *upward* revision of output gaps, even several years after the first releases of data. It can be argued that, with hindsight, the data employed by Galì and Perotti (2003) may be considered as "quasi-real-time" relative to the 1999-2000 expansion, as they are to some extent closer to our "real-time" than to our "revised" ones. In fact, data released in 2002 for the 1999-2000 period have been revised further, and heavily, after 2002. Therefore, it is not surprising that the authors find some evidence of counter-cyclicality, although they made no distinction between *intentional* and *realized* stance. At the same time, and with the perspective of today, the use of revised data (as from the December 2006 OECD Economic Outlook) speaks strongly in favor of pro-cyclicality. In particular, it can be shown that, based on the estimation of a fiscal policy rule as equation 6 (which is virtually equivalent to the one proposed by Galì and Perotti (2003)), the fiscal stance is gauged to be significantly pro-cyclical not only for the whole 19-countries sample, but also considering separately EMU and OECD non-EMU country groups.

[24] The author finds that, in a Barro-Gordon type of framework, and in the presence of 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 tightening 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.

Threshold effects may be also at play as regards the response of fiscal policy to the level of government debt. 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-constant, 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 and whether the way fiscal policy is conducted might be characterized by a multiplicity of regimes.

Hansen's (1999) model is taken as a guideline as it does not rely on an "ad-hoc" 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 $N \rightarrow \infty$) is used to draw valid inference on parameters in different regimes.[25] Hansen's (1999) model has been developed for non-dynamic balanced panels. In Appendix B we recall the key lines of that work, and we propose a two-stage procedure that allows applications to dynamic panel models.

Possible non-linear reactions to cyclical developments and to the debt dynamics are modeled separately, as the approach proposed by Hansen (1999) does not allow to nest in the same framework threshold identification for two different variables. 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. Secondly, and symmetrically, the *capb* is allowed to react non-linearly to the debt indicator, keeping the sensitivity to the output gap invariant. In particular, as concerns the former test, our

[25]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.

approach will allow to select one of the following models

$$\begin{aligned} capb_{i,t|t-1} &= \alpha_i + \rho capb_{i,t-1|t-1} \\ &+\beta x_{i,t-1|t-1} + \theta d_{i,t-1|t-1} + \psi\, emu_{i,t-1|t-1} + \varepsilon_{i,t}; \end{aligned} \tag{10}$$

$$\begin{aligned} capb_{i,t|t-1} &= \alpha_i + \rho capb_{i,t-1|t-1} \\ &+\beta_1 x_{i,t-1|t-1} I(x_{i,t-1|t-1} \leq \gamma_x) \\ &+\beta_2 x_{i,t-1|t-1} I(x_{i,t-1|t-1} > \gamma_x) + \theta d_{i,t-1|t-1} + \psi\, emu_{i,t-1|t-1} + \varepsilon_{i,t}; \end{aligned} \tag{11}$$

$$\begin{aligned} capb_{i,t|t-1} &= \alpha_i + \rho capb_{i,t-1|t-1} \\ &+\beta_1 x_{i,t-1|t-1} I(x_{i,t-1|t-1} \leq \gamma_{1,x}) \\ &+\beta_2 x_{i,t-1|t-1} I(\gamma_{1,x} < x_{i,t-1|t-1} \leq \gamma_{2,x}) \\ &+\beta_3 x_{i,t-1|t-1} I(\gamma_{2,x} < x_{i,t-1|t-1}) + \theta d_{i,t-1|t-1} + \psi\, emu_{i,t-1|t-1} + \varepsilon_{i,t} \end{aligned} \tag{12}$$

where equation (10) is characterized by a single regime (linear fiscal policy rule), equation (11) by two regimes (single threshold rule) and equation (12) by three regimes (double threshold rule) in the way discretionary fiscal policies counteract the economic cycle. In this framework, γ_x is the (unknown) threshold level associated with the output gap in the two-regimes model, $\gamma_{1,x}$ and $\gamma_{2,x}$ the two (unknown) thresholds in the three-regimes model.[26]

Further, possible non-linear reactions to public debt dynamics are tested based on the estimation of the following equations

$$\begin{aligned} capb_{i,t|t-1} &= \alpha_i + \rho capb_{i,t-1|t-1} + \beta x_{i,t-1|t-1} \\ &+\theta d_{i,t-1|t-1} + \psi\, emu_{i,t-1|t-1} + \varepsilon_{i,t}; \end{aligned} \tag{13}$$

$$\begin{aligned} capb_{i,t|t-1} &= \alpha_i + \rho capb_{i,t-1|t-1} + \beta x_{i,t-1|t-1} \\ &+\theta_1 d_{i,t-1|t-1} I(d_{i,t-1|t-1} \leq \gamma_d) \\ &+\theta_2 d_{i,t-1|t-1} I(d_{i,t-1|t-1} > \gamma_d) + \psi\, emu_{i,t-1|t-1} + \varepsilon_{i,t}; \end{aligned} \tag{14}$$

$$\begin{aligned} capb_{i,t|t-1} &= \alpha_i + \rho capb_{i,t-1|t-1} + \beta x_{i,t-1|t-1} \\ &+\theta_1 d_{i,t-1|t-1} I(d_{i,t-1|t-1} \leq \gamma_{d,1}) \\ &+\theta_2 d_{i,t-1|t-1} I(\gamma_{d,1} < d_{i,t-1|t-1} \leq \gamma_{d,2}) \\ &+\theta_3 d_{i,t-1|t-1} I(\gamma_{d,2} < d_{i,t-1|t-1}) + \psi\, emu_{i,t-1|t-1} + \varepsilon_{i,t}; \end{aligned} \tag{15}$$

[26] Hansen's (1999) approach can be virtually extended to models incorporating more than two thresholds. Here at most two thresholds are considered, since it is unlikely that the discretionary behavior of fiscal authorities is characterized by more than three regimes.

where γ_d is the (unknown) threshold level associated with the debt ratio in the single threshold model, $\gamma_{1,d}$ and $\gamma_{2,d}$ the two (unknown) thresholds in the double threshold model.

The notation in equations (10) through (15) refers to the "fully-real-time" case, since only *ex-ante* data are employed. However, tests will be also performed employing revised data for all the variables included given that, 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. Firstly, a "backward-looking" specification (henceforth BL) of the fiscal policy rule is proposed, including lagged observations (revised data) and current-year estimates (real-time data) of output gaps, i.e. $x_{i,t-1}$ and $x_{i,t-1|t-1}$. In addition, as as robustness exercise, tests will be also carried out based on a "forward-looking" specification (henceforth FL), incorporating contemporaneous observations (revised data) and one-year-ahead forecasts (real-time data) of x, i.e. $x_{i,t}$ and $x_{i,t|t-1}$.[27] As in the previous Sections, the sample includes 19 OECD countries, spanning the period 1994 to 2006.

For sake of space, the testing procedure is laid out in Appendix B. Here we just recall that the likelihood ratio statistics F_1 and F_2 are respectively used to test the linear model (i.e. equations (10) and (13)) against the single threshold one (i.e. equations (11) and (14)), and the single threshold model against the double threshold one (i.e. equations (12) and (15)). The likelihood ratio test LR_1 is employed to construct confidence intervals around the estimated $\hat{\gamma}_x$ and $\hat{\gamma}_d$ in single threshold models.

Table 5 reports the F_1 and F_2 statistics, along with their p-values constructed from 5000 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 10% level when a BL rule is considered and at 5% level when the one-year-ahead output gap is used as threshold variable (FL specification). The presence of three regimes is however rejected, as indicated by the p-values associated with the F_2 statistics. This suggests that a single threshold fiscal reaction function (as equation (10)) should be the appropriate model for analyzing the *ex-ante* cyclical stance of fiscal policy.

The estimated $\hat{\gamma}_x$ is in the negative region, respectively equal to -1.2 in the BL case and to -1.8 in the FL one. However, the relatively large 68% and 95% confidence intervals constructed around these estimates generally include positive, and low, values of the output gap. Hence, the threshold might be equally located during mild upturns and slowdowns of the business cycle. Table 6 presents the $\hat{\beta}_1$ and $\hat{\beta}_2$ "regime-dependent" regression estimates suggesting that, broadly in line with results reported in Column 7 and 8 of Table 4, the fiscal policy stance seems neutral when the output gap is below the threshold, whereas strongly and significantly counter-cyclical when it is above it.

[27]The use of contemporaneous data and one-year-ahead forecasts of the output gap in the FL rule may induce an endogeneity problem, since these indicator might be affected by fiscal measures planned for the same year. Although IV estimators should be employed in this case, for simplicity and consistency with the BL exercise we follow Hansen's (1999) model which has been developed based on LS methods.

These findings indicate that asymmetric effects seem to be at play as long as 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 are reported in Table 7. The estimated F_1 statistics strongly point to a rejection of threshold effects, no matter whether real-time or revised data are used in the estimation.[28] It is however interesting to note that, as Figure 4 shows for the "fully-real-time" case and the BL specification, the likelihood ratio LR_1 is minimized when the threshold variable is at 98.3%. 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 (14)), are equal to 0.0116 and 0.0028 respectively. The former is significant at the 5% level while the latter is statistically insignificant. This would imply that, over the last thirteen years, governments have reacted in a more sustainable way to the accumulation of public debt when its level was relatively low.[29]

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:[30]

[28] 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. The hypothesis of multiple regimes is again rejected, both in the real-time and the revised case (results not shown, available from the author). A still unexplored but possibly promising alternative approach could be the one of testing whether governments react non-linearly to implicit liabilities.

[29] As far as EMU countries are concerned, this 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.

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

1. The available observations on real GDP growth rates from year $t_1 = t - T^{start}$ to year $t_2 = t + T^{end}$ are collected;[31]

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.

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 0.11, and significant at the 5% 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 a 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 99% 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

[31] T^{start} and T^{end} define the data span available in each edition of the Economic Outlook. T^{start} is set equal to 16 for vintages up to 2002 and to 13 for the ones from 2003 to 2006. T^{end} is 2 since the Economic Outlook reports the one-year-ahead and two-years-ahead forecasts of real GDP growth rates.

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 output gap for year $t-1$ ($x_{i,t-1|t-1}$) and the $t-1$ current-year output gap (unweighted) averages over all the OECD countries considered, other than 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 close to the benchmark backward looking parameter reported in Column 1. 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 0.37, and significant at a 1% 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 suggested by Buti and van den Noord (2004). The benchmark regression (equation (9)) is augmented by a dummy variable taking value of one in parliamentary election years and zero otherwise.[32] As Column 9 and 10 show, the coefficient associated with this regressor is weakly significant but of the expected negative sign indicating that the occurrence of an election may lead 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 does not seem to be 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 policy rules. We demonstrate that the sign and the size of the bias incurred in estimating a fiscal rule on revised data, when an *ex-ante* relation is under investigation, can be accurately predicted based on empirical covariances among measurement errors, and on other second-order moments. In particular, our findings suggest that the use of updated observations would point to a pro-cyclical fiscal policy stance in industrialized countries over the 1994-2006 period, whereas real-time data indicate that fiscal policy was intended to be counter-cyclical, especially as long as economic expansions are concerned.

Further, formal tests based on Hansen (1999) are performed to explore whether the

[32]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).

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. Multiple regimes 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 doubts on the effectiveness of discretionary fiscal policies to fine tune the business cycle. In fact, albeit the *intentional* stance of the policy seems to be 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 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.

Artis, M. and Marcellino, M. (2001), 'Fiscal Forecasting: The Track Record of the IMF, OECD and EC', *Econometrics Journal* **4**(1), S20–S36.

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

Beetsma, R. and Giuliodori, M. (2008), Fiscal Adjustment to Cyclical Developments in the OECD: an Empirical Analysis Based on Real-Time Data. University of Amsterdam, *mimeo*.

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.

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. (2007), Fiscal Policy in Real Time, CEPII Working Paper No. 07-10, May.

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 Quarterly* **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.

Giuliodori, M. and Beetsma, R. (2008), 'On the Relationship between Fiscal Plans in the European Union: An Empirical Analysis Based on Real-Time Data', *Journal of Comparative Economics (forthcoming)* .

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

Lane, P. R. (2003*b*), 'The Cyclical Behavior of Fiscal Policy: Evidence from the OECD', *Journal of Public Finance* **87**, 2661–2675.

Lane, P. R. (2003*c*), 'The Cyclical Behaviour of Fiscal Policy: Evidence from the OECD', *Journal of Public Economics* **87**(12), 2661–2675.

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.

Romer, C. D. and Romer, D. H. (2007), The Macroeconomic Effects of Tax Changes: Estimates Based on a New Measure of Fiscal Shocks, NBER Working Paper No. 13264.

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

Strauch, R., Hallerberg, M. and von Hagen, J. (2004), Budgetary Forecasts in Europe - the Track Record of Stability and Convergence Programmes, Working Paper Series No. 307, European Central Bank.

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

Taylor, J. B. (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: Revision errors in the dependent variable and inconsistency of Fixed-Effects Least-Squares estimators.

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 of 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

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

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 measurement errors in y and z.[33]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

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

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: $Cov(\nu^{y_i}, \nu^{z_j}) = \sigma_{\nu^{y_i} \nu^{z_j}} = \sigma_{\nu^y \nu^z}$,

[33]Notice that 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.

$\forall i, j \in [1, N]$. Note that we do not introduce any assumption on *auto-correlation* of measurement errors since this does not affect the derivation of the asymptotic bias, as spelled out more in detail below.
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 of 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) through (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

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

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.[34]

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.

[34] Autocorrelation in revision errors in y and z may also be explicitly introduced, for instance by assuming that ν^y and ν^z follow $AR(1)$ processes such as $\nu^y_{i,t} = \rho_{\nu^y}\nu^y_{i,t-1} + \eta^y_t$ and $\nu^z_{i,t} = \rho_{\nu^z}\nu^z_{i,t-1} + \eta^z_t$, where $\rho_{\nu^y} \neq 0$, $\rho_{\nu^z} \neq 0$, $\eta^y_t \sim (0, \sigma^2_{\eta^y})$, $\eta^z_t \sim (0, \sigma^2_{\eta^z})$, $E(\eta^y_t\eta^z_t) = \sigma_{\eta^y\eta^z}$, $E(\eta^y_t\eta^y_{t-k}) = 0$ and $E(\eta^z_t\eta^z_{t-k}) = 0$ for $k = 1\ldots\infty$. In this case, the probability limit of $\hat{\beta}$ would be

$$plim(\hat{\beta}) = \beta + \frac{1}{\sigma^2_{\tilde{z}^*} + \frac{\sigma^2_{\tilde{\eta}^z}}{1-\rho^2_{\nu^z}}}(\frac{\sigma_{\tilde{\eta}^z\tilde{\eta}^y}}{1-\rho_{\nu^z}\rho_{\nu^y}} - \beta\frac{\sigma^2_{\tilde{\eta}^z}}{1-\rho^2_{\nu^z}}). \tag{A.19}$$

Yet, the exercise illustrated in Section 4.2 is based on (A.18) given that the *empirical counterparts* of $\sigma^2_{\tilde{\nu}^z}$,$\sigma_{\tilde{\nu}^z\tilde{\nu}^y}$ and $\sigma^2_{\tilde{\nu}^z}$ in that equation already incorporate possible autocorrelation in revisions. In addition, the use of (A.19) would imply that five more parameters need to be estimated.

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.

B Appendix: A threshold panel regression model

This Appendix illustrates the main building blocks of the panel threshold model by Hansen (1999), before proposing a two-stage procedure for applications to dynamic panels. The single threshold model and the double threshold one will be presented.

B.1 The single threshold model

Let y_{it}, z_{it} and q_{it} be 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

$$y_{i,t} = \alpha_i + \beta z_{it} + \varepsilon_{i,t}, \tag{B.1}$$

or

$$y_{i,t} = \alpha_i + \beta_1 z_{it} I(q_{it} \leq \gamma) + \beta_2 z_{it} I(q_{it} > \gamma) + \varepsilon_{i,t}, \tag{B.2}$$

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 $q_{i,t}$ exceeds a certain threshold γ. The model 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 $\mathbf{q}$ 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 $\hat{\gamma}$.[35] For each q_j (or γ_j), perform the within transformation of equations (B.2).[36] 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,

[35] For each experiment performed in Section 4.4, the trimming parameter η is fixed to 22 to make sure that at least 50 observations ($\simeq N \times (T-1) \times \eta/100$) lie in each regime.

[36] Hereafter γ is used for γ_j.

$$\hat{\gamma} = \underset{\gamma}{\operatorname{argmin}}\, S_1(\gamma). \tag{B.3}$$

2. The H_0 hypothesis of no threshold effects is tested based on the likelihood ratio statistic

$$F_1 = (S_0 - S_1(\hat{\gamma}))/\hat{\sigma}^2, \tag{B.4}$$

where S_0 is the sum of squared errors from the estimation of the linear model (B.1) and $\hat{\sigma}^2$ is the estimated variance of residuals from model (B.2). The asymptotic distribution of F_1 is non-standard, hence 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

$$LR_1(\gamma) = (S_1(\gamma) - S_1(\hat{\gamma}))/\hat{\sigma}^2, \tag{B.5}$$

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

$$LR_1(\gamma) \rightarrow_d \xi \tag{B.6}$$

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

$$P(\xi \leq x) = (1 - exp(-x/2))^2. \tag{B.7}$$

The asymptotic distribution in (B.6) is pivotal, and it may be used to construct asymptotically valid confidence intervals. The distribution function (B.7) 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 3.48 respectively). Finally, the "acceptance region" of confidence level $1-\alpha$ can be derived as that set of values of γ for which $LR_1(\gamma) \leq c(\alpha)$. This can be visually seen by plotting $LR_1(\gamma)$ against a flat line at $c(\alpha)$.[37]

[37] 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.

B.2 The double threshold model

The "true" model may incorporate more than one threshold. In Hansen's (1999) double threshold model, the procedure for estimation and inference on threshold parameters is more cumbersome, albeit intuitively similar, than what shown for the single threshold model. The double threshold regression model reads as

$$y_{i,t} = \alpha_i + \beta_1 z_{it} I(q_{it} \leq \gamma_1) + \beta_2 z_{it} I(\gamma_1 < q_{it} \leq \gamma_2) + \beta_3 z_{it} I(q_{it} > \gamma_2) + \varepsilon_{i,t}, \tag{B.8}$$

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

$$\hat{\gamma}_2^r = \underset{\gamma_2}{\operatorname{argmin}}\, S_2^r(\gamma_2), \tag{B.9}$$

   where

$$S_2^r(\gamma_2) = \begin{cases} S(\hat{\gamma}_1, \gamma_2) & \text{if} \quad \hat{\gamma}_1 < \gamma_2 \\ S(\gamma_2, \hat{\gamma}_1) & \text{if} \quad \gamma_2 < \hat{\gamma}_1 \end{cases} \tag{B.10}$$

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

$$\hat{\gamma}_1^r = \underset{\gamma_1}{\operatorname{argmin}}\, S_1^r(\gamma_1) \tag{B.11}$$

   where

$$S_1^r(\gamma_1) = \begin{cases} S(\gamma_1, \hat{\gamma}_2) & \text{if} \quad \gamma_1 < \hat{\gamma}_2^r \\ S(\hat{\gamma}_2^r, \gamma_1) & \text{if} \quad \hat{\gamma}_2^r < \gamma_1 \end{cases} \tag{B.12}$$

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

$$F_2 = (S_1(\hat{\gamma}_1) - S_2^r(\hat{\gamma}_2^r))/\hat{\sigma}^2. \tag{B.13}$$

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

$$LR_2^r = (S_2^r(\gamma) - S_2^r(\hat{\gamma}_2^r))/\hat{\sigma}^2 \tag{B.14}$$

and

$$LR_1^r = (S_1^r(\gamma) - S_1^r(\hat{\gamma}_1^r))/\hat{\sigma}^2, \tag{B.15}$$

where $S_1^r(\gamma)$ and $S_2^r(\gamma)$ are defined by (B.10) and (B.12). The "no-rejection" regions are the set of values of γ such that $LR_1^r \leq c(\alpha)$ and $LR_2^r \leq c(\alpha)$.

B.3 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 may stem from the inclusion of one (or more) lagged term of the dependent variable. In particular, this holds for models (10) through (15), when *revised-data* are used in estimation (see also equations (6) through (8)). At the same time, using *real-time* variables implies that these models are 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 right-hand-side of the equation. However, these two terms are likely to be highly correlated.

Hence, we propose a two-stage procedure to address this issue. In the following, the procedure for the single threshold model is presented. The same approach applies for the double threshold model (not shown).

Let the dynamic panel model of interest be represented by

$$y_{i,t} = \alpha_i + \rho y_{i,t-1} + \beta z_{it} + \varepsilon_{i,t}, \tag{B.16}$$

where the variable notation is the same above, and where ρ is the autoregressive coefficient associated with the lagged dependent variable $y_{i,t-1}$. The procedure consists of these two steps:

1. In the first-stage, ρ is estimated from equation (B.16) by FE-LS;

2. In the second-stage, the single threshold model is tested against the linear one following the steps laid out in Section B.1, but equations B.1 and B.2 are modified as

$$y_{i,t} = \alpha_i + \hat{\rho} y_{i,t-1} + \beta z_{it} + \varepsilon_{i,t}, \tag{B.17}$$

and

$$y_{i,t} = \alpha_i + \hat{\rho} y_{i,t-1} + \beta_1 z_{it} I(q_{it} \leq \gamma) + \beta_2 z_{it} I(q_{it} > \gamma) + \varepsilon_{i,t}, \tag{B.18}$$

where $\hat{\rho}$ is treated as known.[38] Finally, the likelihood ratio statistic F_1 and LR_1 are used to test H_0 and $\tilde{H}_0$.[39]

[38] In the exercises carried out in Section 4.4, the first-stage regression parameters $\hat{\rho}$ are equal to 0.85 and 0.70 in the real-time and revised data case respectively.

[39] Notice that the two-stage procedure presented here should take into account the uncertainty stemming from the estimation of ρ from the first-stage. As a consequence, confidence intervals from the second-stage should be corrected accordingly. Nevertheless, we treat $\hat{\rho}$ as known. In the empirical application of Section 4.4, the fact that $\hat{\rho}$ is very precisely estimated, both in the BL and the FL case, as shown in Table 4, suggests that the impact of this choice on the final estimates is of second-order importance. In fact, replacing the first-stage (point) estimate with the upper and lower bounds associated with the 95% confidence around $\hat{\rho}_1$ does not affect results considerably (estimates available from the author).

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	0.78	3.83
Greece	0.73	2.94	2.83	13.59
Ireland	1.61	2.12	1.32	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	1.08	6.95
Portugal	1.30	1.56	1.23	4.63
Sweden	1.20	1.72	1.26	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	2.09	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	1.16	1.43	1.15	5.67

Source: Author own calculations based on the December Editions of the OECD Economic Outlook, from Number 56 to 80. Note: As defined in Section (3), $\nu^{x}_{i,t}$,$\nu^{capb}_{i,t}$ 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.

$w =$	(1) $\nu^{x}_{i,t-1}$	(2) $\nu^{capb}_{i,t-1}$	(3) $\nu^{d}_{i,t-1}$
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	0.77	-0.53
Italy	-0.91	0.95	-0.43
France	-0.55	0.74	0.07
Netherlands	-0.68	0.83	-0.15
Portugal	-0.65	0.63	-0.42
Sweden	-0.64	0.84	-0.34
Denmark	-0.67	0.77	-0.43
UK	-0.69	0.85	0.32
Norway	-0.30	0.88	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	0.75	-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 of the *capb* and the revision errors in the current-year estimates (for year $t-1$) of the *capb*,x and d, respectively defined as $\nu^{x}_{i,t-1}$, $\nu^{capb}_{i,t-1}$ and $\nu^{d}_{i,t-1}$ (see Section (3)).

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		
$x_{i,t-1\|t-1}$		0.07	0.35***
		0.78	*3.78*
R^2 (within)	0.025	0.003	0.064
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:	$capb, x, d, emu$		$capb, d, emu$		$capb$		$-$	
Real-time data:	$-$		x		x, d, emu		$capb, x, d, emu$	
Dependent variable	$capb_{i,t}$		$capb_{i,t}$		$capb_{i,t}$		$capb_{i,t\|t-1}$	
Lagged dep.var.	$capb_{i,t-1}$		$capb_{i,t-1}$		$capb_{i,t-1}$		$capb_{i,t-1\|t-1}$	
$\hat{\rho}$	0.70***	0.70***	0.71***	0.71***	0.71***	0.71***	0.85***	0.85***
	16.21	*15.99*	*16.00*	*15.96*	*16.38*	*16.36*	*34.88*	*35.31*
Output gap	$x_{i,t-1}$		$x_{i,t-1\|t-1}$		$x_{i,t-1\|t-1}$		$x_{i,t-1\|t-1}$	
$\hat{\beta}$	-0.15***		-0.07		-0.03		0.07*	
	-3.31		*-1.16*		*-0.41*		*1.78*	
Negative output gap		$x_{i,t-1} \leq 0$		$x_{i,t-1\|t-1} \leq 0$		$x_{i,t-1\|t-1} \leq 0$		$x_{i,t-1\|t-1} \leq 0$
$\hat{\beta}_1$		-0.10		-0.04		0.04		-0.02
		-1.50		*-0.49*		*0.45*		*-0.42*
Positive output gap		$x_{i,t-1} > 0$		$x_{i,t-1\|t-1} > 0$		$x_{i,t-1\|t-1} > 0$		$x_{i,t-1\|t-1} > 0$
$\hat{\beta}_2$		-0.26**		-0.14		-0.15		0.24***
		-2.48		*-1.07*		*-1.19*		*3.23*
Debt	$d_{i,t-1}$		$d_{i,t-1}$		$d_{i,t-1\|t-1}$		$d_{i,t-1\|t-1}$	
$\hat{\theta}$	0.01	0.01	0.02***	0.02***	0.02***	0.02***	0.01	0.01
	1.63	*1.61*	*3.01*	*2.91*	*2.77*	*2.57*	*0.80*	*1.20*
Dummy EMU	$emu_{i,t-1}$		$emu_{i,t-1}$		$emu_{i,t-1\|t-1}$		$emu_{i,t-1\|t-1}$	
$\hat{\psi}$	0.11	0.13	0.29	0.32	0.66***	0.74***	0.47***	0.35**
	0.45	*0.53*	*1.11*	*1.21*	*2.69*	*2.90*	*3.25*	*2.37*
R^2								
Within	0.59	0.59	0.57	0.57	0.58	0.58	0.87	0.87
Between	0.94	0.94	0.86	0.86	0.90	0.90	0.99	0.99
Overall	0.82	0.82	0.77	0.77	0.80	0.80	0.95	0.95
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 (6), (7), (8) and (9), 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	
	BL	FL	BL	FL
	(1)	(2)	(3)	(4)
Test for single threshold				
F_1	7.34	10.44	2.81	0.39
p-value	0.087	0.018	0.493	0.979
critical values:				
10%	6.99	6.29	8.85	6.72
5%	9.00	7.98	11.78	8.56
1%	14.02	11.76	18.99	12.38
Test for double threshold				
F_2	0.12	0.48		
p-value	0.967	0.738		
critical values:				
10%	5.25	3.29		
5%	6.60	4.41		
1%	10.07	6.38		

Notes: BL refers to the "backward looking" specification where $x_{i,t-1|t-1}$ and $x_{i,t-1}$ are used as threshold variables in the real-time and revised data case respectively (Columns 1 and 3). FL refers to the "forward looking" one where $x_{i,t|t-1}$ and $x_{i,t}$ are used (Columns 2 and 4). The autoregressive coefficient from the first-stage regression, $\hat{\rho}_1$, is equal to 0.85 in the real-time data case and to 0.70 in the revised data one. 5000 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	
	BL	FL
Threshold estimates		
$\hat{\gamma}_x$	-1.2	-1.8
68% confidence interval	[-1.5, 1.1]	[-1.9, -0.5]
95% confidence interval	[-2.8, 1.1]	[-2.5, 0.9]
Cyclical sensitivities		
$\hat{\beta}_1$	0.02	-0.06
	0.56	*-1.38*
$\hat{\beta}_2$	0.21***	0.13**
	3.40	*1.99*
Observations in regime 1	96	58
Observations in regime 2	132	170

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	
	BL	FL	BL	FL
	(1)	(2)	(3)	(4)
Test for single threshold				
F_1	9.99	6.60	5.14	3.99
p-value	0.311	0.583	0.676	0.881
critical values:				
10%	15.67	14.68	13.82	14.31
5%	20.29	17.98	15.44	17.01
1%	27.77	25.49	20.57	26.09

Notes 1. The public debt is defined as gross financial liabilities held by the general government.
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: $capb_{i,t|t-1} = \alpha_i + \rho capb_{i,t-1|t-1} + \beta w + \theta d_{i,t-1|t-1} + \psi emu_{i,t-1|t-1} + \varepsilon_{i,t}$

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

	Benchmark		HP Filter		Real-time GDP		Forward looking		Election years	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)
$\hat{\rho}$	0.85***	0.85***	0.86***	0.86***	0.85***	0.85***	0.85***	0.84***	0.85***	0.85***
	34.88	*35.31*	*35.87*	*35.81*	*35.36*	*35.40*	*34.63*	*32.88*	*35.06*	*35.51*
$w =$	$x_{i,t-1\|t-1}$		$x^{hp}_{i,t-1\|t-1}$		$gdp_{i,t-1\|t-1}$		$x_{i,t\|t-1}$		$x_{i,t-1\|t-1}$	
$\hat{\beta}$	0.07*		0.11**		0.11***		0.09*		0.07*	
	1.78		*2.50*		*3.06*		*1.85*		*1.75*	
$\hat{\beta}_1$		-0.02		0.03		-0.01		-0.05		-0.02
		-0.42		*0.35*		*-0.06*		*-0.47*		*-0.42*
$\hat{\beta}_2$		0.24***		0.24**		0.24***		0.37***		0.24***
		3.23		*2.35*		*3.12*		*3.17*		*3.25*
$\hat{\theta}$	0.01	0.01	-0.01	-0.01	-0.01	0.01	0.01	0.01	0.01	0.01
	0.80	*1.20*	*-0.35*	*-0.25*	*-0.11*	*0.01*	*0.75*	*1.16*	*0.17*	*0.49*
$\hat{\psi}$	0.47***	0.35**	0.41***	0.41***	0.45***	0.45***	0.50***	0.35**	0.47***	0.35**
	3.25	*2.37*	*2.96*	*2.94*	*3.24*	*3.24*	*3.32*	*2.23*	*3.26*	*2.37*
Elections									-0.16	-0.16*
									-1.61	*-1.68*
Method	FE-LS	FE-LS	FE-LS	FE-LS	FE-LS	FE-LS	IV	IV	FE-LS	FE-LS
R^2										
Within	0.87	0.87	0.87	0.87	0.87	0.87	-	-	0.87	0.87
Between	0.99	0.99	0.99	0.99	0.99	0.99	-	-	0.99	0.99
Overall	0.95	0.95	0.95	0.95	0.95	0.95	0.86	0.86	0.95	0.95
Sargan-p	-	-	-	-	-	-	0.18	0.11	-	-
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 $t-1$) output gap (positive and negative output gaps are used in the regression of Column 8) 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.

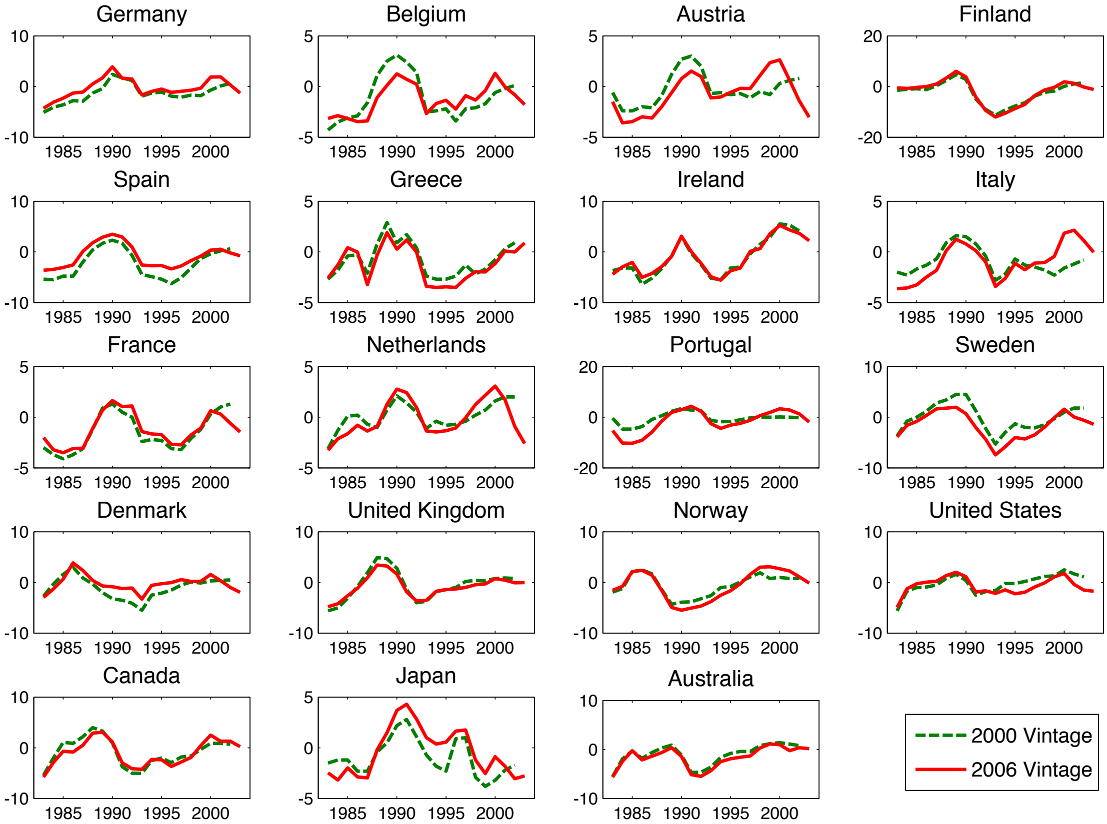

Figure 1: Output gaps for some OECD countries from two data vintages. Data sources: OECD Economic Outlook No. 68 (December 2000) and No. 80 (December 2006).

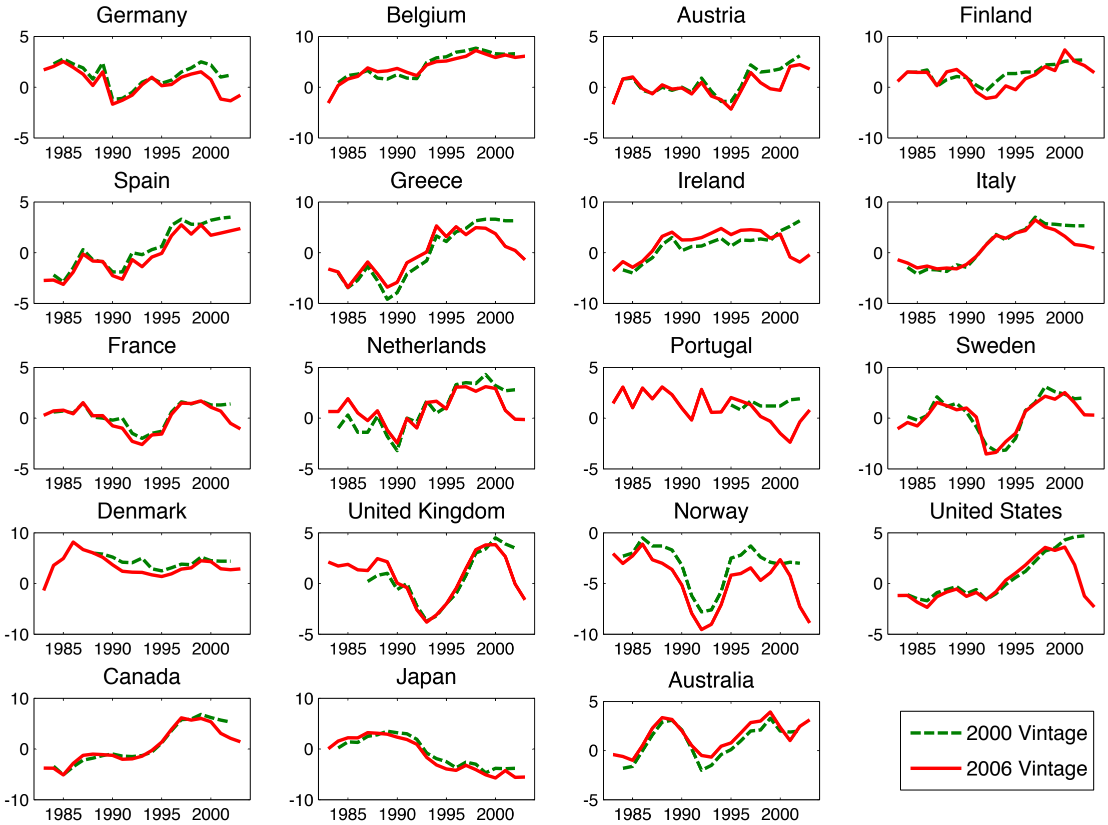

Figure 2: Cyclically-adjusted primary balances as percentage of potential GDP for some OECD countries from two data vintages. Data sources: OECD Economic Outlook No. 68 (December 2000) and No. 80 (December 2006).

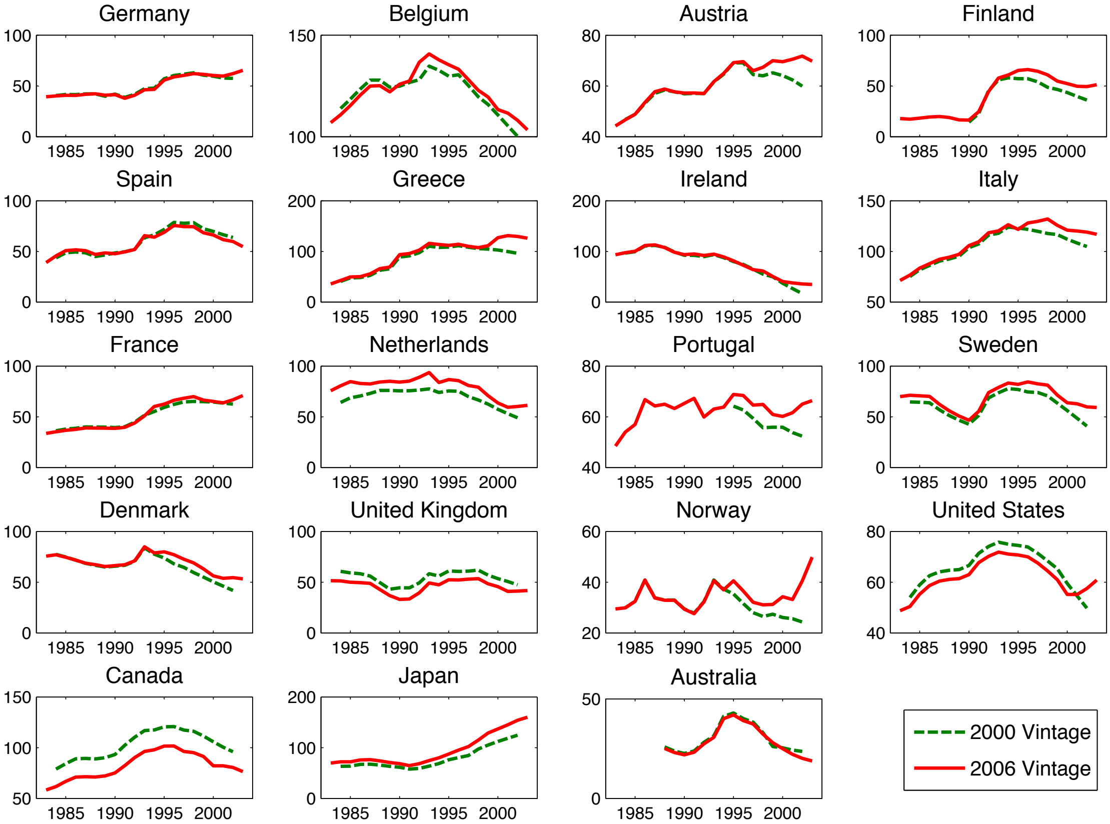

Figure 3: General government gross financial liabilities as percentage of GDP for some OECD countries from two data vintages. Data sources: OECD Economic Outlook No. 68 (December 2000) and No. 80 (December 2006).

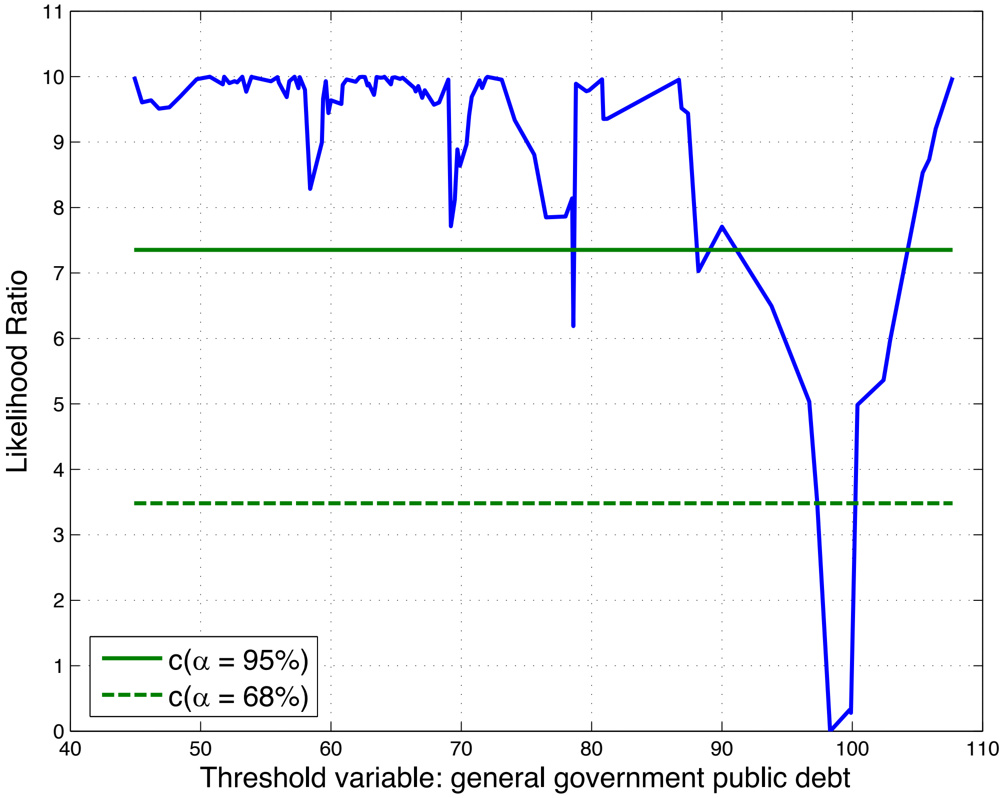

Figure 4: Single threshold estimation. The graph refers to the case BL specification, where the threshold variable is the general government public debt as percentage of GDP and real-time data are used (see equation (14)). Blue lines: likelihood ratio statistic LR_1 (equation (B.5)); green solid lines: 95% critical value (10.59); green dotted line: 68% critical value (7.35).

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