Universidade do Minho

Documentos de Trabalho

Working Paper Series

"Assessing Fiscal Sustainability Subject to Policy Changes: A Markov Switching Cointegration Approach"

Vasco J. Gabriel

Pataaree Sangduan

NIPE WP 9 / 2009

NÚCLEO DE INVESTIGAÇÃO EM POLÍTICAS ECONÓMICAS

UNIVERSIDADE DO MINHO

"Assessing Fiscal Sustainability Subject to Policy Changes: A Markov Switching Cointegration Approach"

Vasco J. Gabriel
Pataaree Sangduan

NIPE* WP 9/ 2009

URL:
http://www.eeg.uminho.pt/economia/nipe

* NIPE – *Núcleo de Investigação em Políticas Económicas* – is supported by the Portuguese Foundation for Science and Technology through the *Programa Operacional Ciência, Teconologia e Inovação* (POCI 2010) of the *Quadro Comunitário de Apoio III*, which is financed by FEDER and Portuguese funds.

Assessing Fiscal Sustainability Subject to Policy Changes: a Markov Switching Cointegration Approach

Vasco J. Gabriel*

Department of Economics, University of Surrey, UK and NIPE-UM

Pataaree Sangduan

Bureau of the Budget, Thailand

January 2009

Abstract

We propose a Markov switching cointegration approach to assess long run fiscal sustainability. This method allow us to simultaneously: 1) test for cointegration in the presence of significant fiscal policy changes; 2) assess the type of fiscal regime (whether 'strongly'/'weakly' sustainable or unsustainable) that a country experienced at a given period and 3) analyse the timing of the transition between the estimated regime types. Given its flexibility, our approach enable us to uncover a richer and more complex dynamics in the analysis of fiscal sustainability, which standard linear cointegration methods fail to capture.

JEL Classification: C22; E62; H60

1 Introduction

A stable and sustainable long-term relationship between government expenditures and revenues is a key requirement for macroeconomic stability. Given its relevance, this issue has attracted a great deal of attention, with particular emphasis on testing empirically whether or not a given country's fiscal stance is sustainable. There is, however, a contradiction between the predictions of empirical models, which point to a significant degree of unsustainability across different

*Corresponding author. Address: Department of Economics, University of Surrey, Guildford, Surrey, GU2 7XH, UK. Email: v.gabriel@surrey.ac.uk. Tel: + 44 1483 682769. Fax: + 44 1483 689548.

countries, and the relative scarcity of episodes of full-scale defaults. It is, therefore, of great importance to reassess empirical methodologies dealing with the analysis of fiscal sustainability.

In this paper, we argue that regime changes are a pervasive feature in the empirical analysis of fiscal sustainability and that, once an appropriate testing method is put to use, the paradoxical findings of earlier literature virtually disappear. Indeed, failure to detect and account for parameter shifts is a serious form of misspecification, therefore affecting inference and leading to poor forecasting performances. This is especially relevant for cointegration analysis, since it normally involves long time spans of data, which, consequently, are likely to display structural breaks.

The possibility of regime changes affecting the results of empirical tests has been recognized early on, namely by Wilcox (1989) and Hakkio and Rush (1991). These authors split their sample (of US data) at exogenously chosen break dates, but this may be problematic, as subsequent tests may have their power affected if the chosen date does not correspond to the true one. The situation when there is no *a priori* information requires a particular type of analysis, so the adopted solution has been to endogeneize the break point selection in the testing problem, maintaining the inference valid. Thus, Haug (1995) - using the tests proposed by Hansen (1992), Quintos (1995) - allowing for changes in the cointegration rank, and Martin (2000) - employing a Bayesian approach to detect multiple breaks, use procedures to endogenously select the break point in the sample (arriving at different conclusions regarding the existence of structural breaks, however).

In this paper, we pursue a different route. We initially test whether or not the long run relationship has been subject to structural breaks. We do so by employing the tests proposed by Gregory and Hansen (1996), as in Baharumshah and Lau (2007), but in addition we extend these tests to the case of a cointegrating relationship without constant term. We then apply the method proposed by Gabriel, Psaradakis and Sola (2002) to investigate cointegration subject to possible changes in regime. As in Hall, Psaradakis and Sola (1997), this formulation is based on the assumption that cointegration regimes are governed by an unobserved Markov chain process. According to Gabriel *et al.* (2002), testing for cointegration may be carried out by means of standard residual-based tests, using the standardized residuals obtained from Markov switching estimation[1].

Markov switching models have been extensively (and successfully) used to characterize and account for regime changes that typically occur in economic and financial time series, such as

[1]See Davies (2006) and Alexandre *et al.* (2007) for applications of this technique.

GDP, stock prices, interest rates, inflation rates, or exchange rates, for example (see Kim and Nelson, 1999 for a survey). Given their flexibility, it would be natural to extend their use to model changes in long run relationships. Hall *et al.* (1997) and Krolzig (1997), for example, illustrate the usefulness of such a specification by analysing the Japanese consumption function and co-movements in international business cycles, respectively. The Markov switching cointegration approach is also related, from a methodological point of view, with the work of Hansen (2003), as this author generalizes Johansen's cointegrated VAR model by allowing for structural breaks.

This approach offers a number of advantages. First, one can resort to the usual asymptotic critical values for residual-based tests, as the finite-sample distributions of the standardized residuals are well approximated by the usual asymptotic distributions. Secondly, previous papers either consider a single, deterministic break or assume that the break points are known when cointegration is being tested. Instead, a Markov switching approach more flexible, as it allows for an unspecified number of breaks, of unknown location. Moreover, information on the timing of the breaks is a natural by-product of estimation. Thirdly, one can also assume changes in the variance of the long run relationship. Furthermore, testing for cointegration arises naturally from the estimation step, since only standard cointegration testing procedures are used. Specifying long run relationships in this way encompasses a number of empirically plausible and economically relevant models, including the case of a single permanent regime change, as discussed below.

The paper is structured as follows. The next section presents the analytical framework for testing fiscal sustainability. Section 3 discusses preliminary empirical analysis, followed by the application of the cointegration tests of Gregory and Hansen (1996) in section 4. We then use the Markov switching approach outlined above to test for fiscal sustainability. A final section concludes.

2 Theoretical framework

Tests of fiscal sustainability are commonly based on the government's intertemporal budget constraint (IBC) in its present value form. Given (in real terms) government expenditures G, revenues R, public debt B and the interest rate i, the government's one-period budget constraint is written as

$$G_t + (1+i)B_{t-1} = R_t + B_t.$$

A similar condition holds for periods $t+1$, $t+2, ...$ with forward substitution yielding the IBC

$$B_t = \sum_{j=0}^{\infty} \Pi_{k=1}^{j} (1 + i_{t+k})^{-1} (R_{t+j} - G_{t+j}) + \lim_{j \to \infty} \Pi_{k=1} (1 + i_{t+k})^{-1} B_{t+j}, \tag{1}$$

implying that current government debt B_t must be financed by the present value of the future primary surpluses[2]. Assuming that interest rates are stationary, the above expression can be conveniently rewritten for empirical purposes as

$$GG_t - R_t = \sum_{j=0}^{\infty} \delta^{j-1} (\Delta R_{t+j} - \Delta GG_{t+j} + i \Delta B_{t+j-1}) \tag{2}$$

where GG_t is now government expenditures inclusive of interest payments with discount factor $\delta = (1 + r)^{-1}$.

Given that the variables GG_t and R_t usually display non-stationary behaviour, this provides a statistical framework for testing sustainability. Indeed, fiscal sustainability implies that revenues and expenditures must be cointegrated, if each are $I(1)$ processes. In practice, this amounts to estimate the generic regression equation

$$R_t = a + bGG_t + u_t \tag{3}$$

and, depending on the cointegration vector $[1, -b]$ obtained, we may have three possible scenarios for sustainability analysis:

- 'Strong' sustainability, if and only if the $I(1)$ processes R_t and GG_t are cointegrated and $b = 1$
- 'Weak' sustainability in the case where R_t and GG_t are cointegrated, but $0 < b < 1$: a smaller than 1 long-run elasticity of revenue relative to expenditure may be an incentive for debt default.
- Unsustainability, when $b \leq 0$, implying that deficits are being accumulated at a rate greater than the growth rate in the economy and the IBC is therefore violated.

Thus, the common procedure in the literature[3] is to apply cointegration tests to (3) (see Haug, 1991, Hakkio and Rush, 1991 and Ahmed and Rogers, 1995, for example). However, evidence of sustainability is mixed. One potential shortcoming of the cointegration methodology outlined above is that the relationship is assumed to be invariant. Given that fiscal policy is

[2] Ruling out Ponzi games and therefore the second, asymptotic term should converge to 0.

[3] Another possibility it to test for a unit root in B_t (see Hamilton and Flavin, 1986 and Wilcox, 1989).

often subject to abrupt changes, motivated by political or economic reasons, this may lead to periods of sustained deficits, which may have important implications for the statistical analysis of fiscal sustainability, resulting in apparent global unsustainability, as shown in Haug (1995) and Quintos (1995).

Therefore, in what follows, we propose an alternative methodology to deal with potential changes in fiscal stances, by assuming that the long run relationship between government revenues and expenditures is subject to Markov-type shifts. We start by looking at the results of standard cointegration analysis, then testing for sustainability allowing for deterministic shifts and finally considering Markov-switching cointegration.

3 Preliminary empirical results

For illustration purposes, we test the fiscal sustainability of a variety of developed and developing countries, namely the Bahamas, Finland, France, South Africa, Thailand and the United States. We use quarterly data for the relevant variables (in real terms), spanning from 1975 to 2004 and collected from the International Financial Statistics database. While initial studies have focused on developed economies (see Payne, 1997, for example), increasing attention has been devoted to the fiscal stance of developing countries (see Kalyoncu, 2005, Baharumshah and Lau, 2007 and Payne, Mohammadi and Cak, 2008, for example). Empirical evidence is ambiguous, suggesting that the case of 'weak' sustainability is very common, particularly for developing economies.

Preliminary unit root tests have largely confirmed that government revenues and expenditures for all countries appear to follow $I(1)$ processes (results available upon request). This suggests that cointegration is the appropriate framework to assess the sustainability of these fiscal regimes. Thus, we explore cointegration inference involving these two variables, by estimating the cointegration regression (3) and testing whether $b = 1$ or $0 < b < 1$.

We employ a residual-based approach to testing cointegration, i.e., we first estimate (3) and then ascertain whether the estimated equilibrium errors are stationary or not, by means of ADF-type and Phillips-Perron-type tests, which are also known as Augmented-Engle-Granger (AEG) and Phillips-Ouliaris (PO) cointegration test. There are no efficiency losses in pursuing a single-equation route when compared to the multi-equation method of Johansen (1988), as we are studying a bivariate relationship with potentially a single cointegration vector. Thus, for conciseness, we consider the standard OLS estimator of b, as well as the dynamic OLS (DOLS) estimator of Stock and Watson (1993), which augments the cointegrating regression with p lags and leads[4] of the differenced explanatory variable, in order to correct for second-order biases usually associated with the simple OLS estimator.

< Insert Table 1 here >

We observe from Table 1 that, in general, the OLS estimates tend to be further away from 1 that the corresponding DOLS estimates ($\hat{b}_{DOLS}$). Considering the estimates alone, this would imply that the Bahamas, Finland and France would be classified as 'weakly' sustainable, with the remaining countries to be considered 'strongly' sustainable. If one looks at the DOLS results, however, all countries display estimates very close to the 'strong' sustainability benchmark, with the exception of Finland, with $\hat{b} = 0.824$.

Note, however, that this analysis is conditional on the existence of cointegration between expenditures and revenues. Looking at the residual-based tests with OLS residuals, one would conclude that, according to the AEG test, Thailand, Finland and the USA would fail to meet the sustainability criteria, given that the statistic fails to reject the null of no cointegration. Interestingly, however, the PO test indicates that only the US would not be sustainable.

If we consider instead tests based on the DOLS estimator, the AEG would point to unsustainability for all countries with the exception of France. The PO test, on the other hand, would add South Africa and the Bahamas to the latter. Therefore, a contradiction seems to emerge: by employing a theoretically more appealing estimator, the case for sustainability is weakened, although the point estimates suggest that the cointegration vector is indeed $[1, -1]$.

It appears that the results of conventional methodologies tend to penalize the sustainability hypothesis, even when the estimated b is close to 1. This could be explained by the fact that regime shifts tests may lead to loss of power of unit root and cointegration tests and, hence, the null hypothesis of cointegration is rejected less often than it should. We next test for sustainability allowing for regime shifts.

4 Testing for sustainability allowing for regime shifts

Gregory and Hansen (1996), generalized the standard cointegration tests by considering an alternative hypothesis in which the cointegration vector may suffer a regime shift at an unknown

[4]We determine p by testing down the significance of the extra leads and lags, starting from $p = 4$.

timing. They analyzed models that accommodate under the alternative hypothesis of cointegration the possibility of changes in parameters, namely an intercept shift model (C), a model with an intercept shift plus trend (C/T), a "regime shift" model (C/S) where both the constant and slope parameters change, as well as a regime shift model where a trend shift is added $(C/S/T)$, see paper for details.

As with the previous tests, Gregory and Hansen (1996) tests are residual-based cointegration procedures that evaluate if the error term is $I(1)$ under the null of no cointegration. In this framework, however, since the change point or its occurrence are unknown, the testing procedures involve computing the usual statistics (AEG, Z_t and Z_α) for all possible break points $\tau \in J$ and then selecting the smallest value obtained, since it will potentially present greater evidence against the null hypothesis of no cointegration[5].

Nevertheless, as pointed out by the authors, these tests possess power against other alternatives, namely "stable" cointegration. Hence, a rejection of the null hypothesis does not necessarily imply changes in the cointegration vector, since an invariant relationship might be the cause of the rejection. Also, note that the smallest value of the statistic, if leading to a rejection, can provide an idea of where a shift might have occurred.

These test statistics have non-standard limiting distributions with no closed form and, therefore, critical values were obtained by resorting to simulation methods. In this section, we examine types of structural breaks that were not previously tabulated by Gregory and Hansen (1996), which are the change in slope with stable intercept,

$$R_t = \mu + \beta_1' GG_t + \beta_2' GG_t D_t + u_t, \tag{S}$$

as well as a model with change in slope and no constant term,

$$R_t = \beta_1' GG_t + \beta_2' GG_t D_t + u_t. \tag{S_{nc}}$$

The vector (R_t, GG_t) is assumed to be of $I(1)$ variables, u_t should be a stationary disturbance and D_t is a dummy variable of the type

$$D_t = \begin{cases} 0, \text{ if } t > [T\tau] \\ 1, \text{ if } t \leq [T\tau] \,. \end{cases} \tag{4}$$

[.] denoting the integer part operator. These models are of interest to the empirical analysis of fiscal sustainability, as there can be cases where fiscal regimes shift between 'strong' and 'weak'

[5] τ denotes the unknown relative timing of the break point and the trimming region is $J = (0.15, 0.85)$, following Gregory and Hansen (1996).

sustainability, as in (S). Furthermore, we also consider the theory-consistent case of no intercept, as the IBC implicitly assumes zero-deficits. The justification for the use of an intercept term in empirical studies is mainly computational, as it ensures that residuals have zero-mean.

For proper comparison, and following Gregory and Hansen (1996, p. 110), we obtained critical values for these types of shifts, with a single regressor, using the same response surface: with 10 000 replications for sample dimensions $T = 50, 100, 150, 200, 250$ and 300, critical values at the p percent level are obtained and then the regression

$$C(p, T) = \psi_0 + \psi_1 T^{-1} + \text{error},$$

is run. The critical values at the 5% significance level for the (S) model are -4.685 (*GH-AEG* and *GH-*Z_t tests) and -39.172 (*GH-*Z_α test). For the (S_{nc}) model, the critical values are -4.192 for the *GH-AEG* and *GH-*Z_t tests, and -30.322 for the *GH-*Z_α test, respectively. The critical values for the (C/S) model, to be used below, are -4.95 and -47.04.

We focus our attention on tests of variants (C/S), (S) and (S_{nc}), as these are the most relevant for our application. Indeed, there is no theoretical reason to expect the cointegrating relationship to have a deterministic trend and the parameter of interest in our case is β. The results are shown in Table 2.

< Insert Table 2 here >

Overall, we find that the null of no cointegration is rejected by the majority of tests, for the three model variants and across the six countries. We notice that the *GH-AEG* test rejects the null less often, while the Z-type tests almost always reject the null. The variant (S_{nc}) is not rejected for Thailand, while in the case of Finland, only the *GH-*Z_t is able to reject. Otherwise, for every country, model variants (C/S) and (S) have their nulls rejected by at least two test statistics. The general conclusion seems to point to the existence of a long-run equilibrium between government expenditures and revenues, but one that appears to have been subject to regime shifts.

In the table, we also report the dates corresponding to the smallest value attained by each statistic. As mentioned above, we can use this as an informal way of dating potential regime shifts. It is interesting to note that for several statistics, the minimum point appears around the Asian crisis of 1997-1998, with the first quarter of 1998 the most often identified date. These results seem to be consistent with the stylized fact observed in many countries, which have experienced fiscal difficulties following the Asian turmoil. This seems to be the case for France, Finland, South Africa and Thailand. In the case of the Bahamas, breaks are also informally identified in the early 90s, while for the US, 1995 appears to signal a shift in the fiscal regime. This coincides with the start of the surplus years of the Clinton Administration.

Thus, it seems appropriate to try to model fiscal sustainability as potentially being subject to regime shifts. However, the tests of this section assume that shifts occur in a deterministic fashion, which is perhaps not very realistic. Also, the timing of the shifts may not be accurate, as the above procedures will signal the largest break in the series. For the sample period considered, it is likely that more than one break as occurred, as it contain years of fiscal difficulties, which at same point appear to have been resolved. Thus, the Gregory-Hansen tests, while being informative in terms of cointegration inference, do not offer a convenient framework to model long run relationships subject to regime changes. A possible way of allowing for stochastic shifts is to use a Markov switching approach, as explained in the next section.

5 Fiscal sustainability under Markov Switching regime changes

In this section, and following Hall *et al.* (1997), we propose to use a more general type of cointegration, where the cointegrating vector is allowed to undergo occasional changes, which may be the result of sudden changes in policy, economic conditions, technology or institutions. In order to describe the long run relationship between revenue and expenditures, we will use the following model

$$R_t = (\alpha_1 + \alpha_2 s_t) + (\beta_1 + \beta_2 s_t)GG_t + (\omega_1 + \omega_2 s_t)u_t \tag{5}$$

where s_t is the discrete-valued latent random variable indicating the regime operative at time t and u_t is a stationary and ergodic random disturbance with mean 0 and unit variance. The variable s_t is assumed to follow a homogeneous first-order Markov chain with state space $\{1, 2\}$ and transition probabilities $p = \Pr(s_t = 2|s_{t-1} = 2)$, $q = \Pr(s_t = 1|s_{t-1} = 1)$. Accordingly, the cointegrating vector will have two regimes defined by s_t, $\{(\alpha_1, \beta_1), (\alpha_2, \beta_2)\}$, while $\omega_{s_t} = \{\omega_1, \omega_2\}$, so that we allow the variance of the long run relationship to change stochastically as well, thus capturing potential low and high volatility regimes. Note that this is a generalization of the Gregory and Hansen (1996) models discussed above, in that we allow the shifts to be stochastic as opposed to deterministic. In addition, they can occur more than once and the variance is allowed to change. In fact, the Gregory-Hansen models correspond to the case where one of the regimes is 'absorbing', that is, the staying probability of one of the regimes is 1.

As suggested by Gabriel *et al.* (2002), one can test for cointegration simply by resorting to the standard residual-based procedures, but using instead the standardized residuals obtained from the estimation of the Markov switching cointegrating model. These residuals are computed as

$$e_t = \frac{R_t - [(\alpha_1 + \beta_1 GG_t)\Pr(s_t = 1|I_t) + (\alpha_2 + \beta_2 GG_t)\Pr(s_t = 2|I_t)]}{[\hat{\omega}_1^2 \Pr(s_t = 1|I_t) + \hat{\omega}_2^2 \Pr(s_t = 2|I_t)]^{1/2}}, \tag{6}$$

where $\Pr(s_t = i|I_t)$, $i = 1, 2$, are the filter probabilities from the Markov switching estimation. If more than one shift has occurred, the usual residuals will reflect this by appearing to be non-stationary and thus cointegration may not be detected. By allowing for an unspecified number of regime changes in the estimation step, the standardized residuals will be free of unusual observations due to breaks, and therefore will replicate the stationary behaviour of the true errors.

< Insert Figure 1 here >

Estimation of (5) is carried out by maximum likelihood[6]. We start by analysing the stationarity of the standardized residuals, computed as in (6) and contrasted in Figure 1 with the residuals obtained from the simple linear regression (3). We can see that the former appear to be more stable and, hence, stationary. Indeed, inspection of the DOLS residuals obtained in section 3 reveals that some countries, in particular Finland, Thailand and the USA, experienced periods of persistent, but temporally circumscribed, deviation from their average time series path (deficits in the case of the first two countries, surpluses in the case of the latter). Given that a linear approach will not be able to model these deviations, the (D)OLS residuals appear to be non-stationary, as suggested by the cointegration tests in section 3. The Markov switching approach discussed here allow us more flexibility in incorporating the regime changes and thus reflecting them in the inference step.

The stationarity of the standardized residuals can be assessed by residual-based cointegration tests. Gabriel *et al.* (2002) show that the asymptotic distributions of these tests provide a good approximation when standardized Markov switching residuals are used. The critical values

[6] Estimates were obtained with a numerical optimization procedure using the BFGS algorithm. The corresponding asymptotic standard errors in Table 4 are heteroskedasticity and autocorrelation consistent (HAC), computed with the prewhitened quadratic spectral kernel and data-dependent bandwidth, as recommended by Andrews and Monahan (1992).

obtained by McKinnon (1991) for these tests are −3.9001, −3.3377 and −3.0462 for 1%, 5% and 10% significance levels, respectively.

< Insert Table 3 here >

Table 3 reports the results of the AEG and PO tests for each country. As before, the lag length for the AEG test was automatically selected based on the SIC procedure, while the bandwidth for the Phillips-Ouliaris test is also data-dependent, based on a Bartlett kernel (results do not change if other intermediate procedures are used instead). We can see that the null hypothesis of cointegration is always comfortably rejected at the 1% significance level for all countries. Thus, comparing with the results of section 3, the conclusion of unsustainability in the case of Finland, Thailand and the USA suggested by the standard cointegration tests is now overturned by the Markov switching-based tests. Indeed, it seems that when we account for regime changes, fiscal sustainability receives stronger empirical support.

< Insert Table 4 here >

Another advantage of this framework is that one can interpret changes in the cointegration vector as shifts in fiscal regimes. Table 4 displays the parameter estimates arising from estimation of (5). We compute, in the last two columns, the difference between the AIC and SIC for the Markov switching and the simple linear model, so that a negative number favours the non-linear specification, which is the case for all countries[7]. In addition, the estimated regimes appear to be quite persistent, with estimated probabilities well above 0.9. We also note that, with the exception of Thailand and South Africa, the variance appears to be the same across regimes.

< Insert Figure 2 here >

Taking each country in turn, we observe that for the Bahamas regime 1 corresponds to a period of 'strong' sustainability, as the estimate of β is not significantly different from 1. However, in state 2, the fiscal regime appears to be unsustainable, as β becomes close (statistically

[7] See Psaradakis, Sola and Spagnolo (2004) on the usefulness of model selection criteria such as the AIC and SIC to detect Markov switching behaviour.

speaking) to 1. Looking at the corresponding panel in Figure 2, which displays the filtered probabilities of regime 1, we can see that periods of instability have mainly occurred in 1984-86 and in 1988-1995, after which sustainability seems to have been resumed.

The same pattern is present in the case of France, in which β drops considerably from 0.9664 to a value statistically close to, or below, 0. The period of fiscal unsustainability coincides with the 'Euro-sclerotic', low growth and high unemployment years of the 90s, which lead to increased pressure on government spending.

In the case of Finland, again regime 1 corresponds to 'strong' sustainability (with $\beta =$ 1.0365), while regime 2 sees a shift to a 'weakly' sustainable regime, since the β becomes lower, but different from 1. The filtered probabilities in the corresponding panel of Figure 2 identify the shift to regime 2 around 1991, which then lasts until 1998. This corresponds to a recessionary period following the collapse and dismantling of trade with the Soviet Union, accompanied by an increase in interest rates in Europe, which drove the currency up (under a pegged exchange rate) and later on to a banking crisis. The economy started to recover from the recession in late 1993 and the deficits started to gradually decline after the fiscal consolidation programme initiated in 1995.

A similar pattern of switches between 'strong' and 'weak' sustainability is apparent for South Africa. For Thailand, the increase in volatility of the fiscal regime since 1988 drives the procedure to identify transitions which seem to coincide with switches in the variance rather than in the strength of fiscal sustainability. The same effect seems to be present in the case of the US, although to a lesser extent. Both countries seem to be well within the 'weak' sustainability case, but the filtered probabilities correctly identify, for the US case, the troublesome periods of the 70s and early 90s already identified in the literature (Hakkio and Rush, 1991, Quintos, 1995 and Martin, 2000), corresponding to a tighter monetary policy and the Reagan Administration tax cuts policies. This was followed by a period of smaller deficits in the late 80s and then the accumulation of surpluses during the late Clinton Administration years.

Thus, with this approach we are able to uncover a richer and more complex dynamics in the analysis of fiscal sustainability, which standard linear cointegration methods fail to capture. Indeed, the results in section 3 pointed to sustainability problems for some countries. However, we have shown that unsustainable fiscal stances have been, to a great extent, temporally circumscribed. This suggests that failure to account for multiple regime changes may affect empirical tests of fiscal sustainability.

6 Conclusion

There is ample evidence in the literature that policy changes or sudden shifts in economic conditions may have a substantial impact on the dynamics of fiscal deficits. In statistical terms, if these changes are left unaccounted for, then a policy that is sustainable overall might appear to be unsustainable. Indeed, if one uses conventional residual-based procedures, structural breaks induce an increase in the residuals autocorrelation which may induce an near-unit root type of behaviour.

By employing cointegration tests specifically designed to take potential regime shifts into account, we have shown that structural breaks seems to be pervasive in tests of fiscal sustainability. However, given that economies may experience periods of limited duration of fiscal stress, modelling this by only estimating breakpoints as in Haug (1995) and Quintos (1995) seems to carry little information other than potential timings of changes. Therefore, we proposed an alternative and more flexible methodology to deal with potential changes in fiscal regimes. By employing a Markov switching specification of the long run relationship between revenues and expenditures, as in Hall *et al.* (1997), we are able to simultaneously: 1) test for cointegration using Gabriel's *et al.* (2002) procedure; 2) assess the type of fiscal regime (whether 'strongly'/'weakly' sustainable or unsustainable) that a country experienced at a given period and 3) analyse the timing of the transition between the estimated regime types.

An alternative to the results presented here would be model the primary surplus/deficit series as a Markov switching process. In principle, similar conclusions would emerge in the case of economies that switch between sustainability and unsustainability. However, it should be noticed that this approach imposes the cointegrating vector $[1, -1]$ throughout and therefore does not allow to distinguish the cases where switches occur between 'strong' and 'weak' sustainability, as the formulation proposed here does.

There is scope for further refinements. As mentioned before, further insight may be gained if one allows the variance of the long run relationship to follow an independent Markov chain. This means that the model can be rewritten as a four-state Markov switching model (see Hall *et al.* 1997 for an example) and then test the hypothesis of whether changes in the variance and in the mean follow the same unobserved latent process. However, this is beyond the scope of the present work, which aims at illustrating the usefulness of this Markov switching approach. We leave this for future research.

References

[1] Ahmed, S. and Rogers, J. H.(1995), Government budget deficits and trade deficits Are present value constraints satisfied in long-term data?, *Journal of Monetary Economics,* 36, 351-374.

[2] Alexandre, F., Bacao, P and Gabriel, V. J. (2007), Volatility in asset prices and long-run wealth effect estimates, *Economic Modelling*, 24, 1048-1064.

[3] Andrews, D. W. K. and Monahan, J. C. (1992), An Improved Heteroskedasticity and Autocorrelation Consistent Covariance Matrix Estimator, *Econometrica,* 60, 953-966.

[4] Baharumshah, A. Z. and Lau, E. (2007), Regime changes and the sustainability of fiscal imbalance in East Asian countries, *Economic Modelling*, 24, 878-894.

[5] Davies, A. (2006), Testing for international equity market integration using regime switching cointegration techniques, *Review of Financial Economics*, 15, 305-321.

[6] Gabriel, V. J., Psaradakis, Z. and Sola, M. (2002), A simple method of testing for cointegration subject to multiple regime change, *Economics Letters*, 76, 213-221.

[7] Gregory, A. W. and Hansen, B. E. (1996), Residual-based Tests for Cointegration in Models with Regime Shifts, *Journal of Econometrics*, 70, 99-126.

[8] Hakkio, C. S. and Rush, M. (1991), Cointegration and Government Borrowing Constraints: Evidence for the United States, *Journal of Business and Economic Statistics*, 9, 429-445.

[9] Hall, S. G., Psaradakis, Z. and Sola, M. (1997), Cointegration and changes in regime: the Japanese consumption function, *Journal of Applied Econometrics*, 12, 151-168.

[10] Hamilton J. D. and Flavin M. A. (1986), On the Limitations of Government Borrowing. A framework for Testing, *American Economic Review*, 76, 808-819.

[11] Hansen, B. E. (1992), Tests for Parameter Instability in Regressions with I(1) Processes, *Journal of Business and Economic Statistics*, 10, 321-335.

[12] Hansen, P. R. (2003), Structural Breaks in the Cointegrated Vector Autoregressive Model, *Journal of Econometrics*, 114, 261-295.

[13] Haug, A. A. (1991), Cointegration and Government Borrowing Constraints: Evidence for the United States, *Journal of Business and Economic Statistics*, 9, 97-101.

[14] Haug, A. A. (1995), Has Federal Budget Deficit Policy Changed in Recent Year?, *Economic Inquiry*, 33, 104-118.

[15] Johansen, S. (1988), Statistical Analysis of Cointegration Vectors, *Journal of Economic Dynamics and Control*, 12, 231-254.

[16] Kalyoncu, H. (2005), Fiscal policy sustainability: test of intertemporal borrowing constraints, *Applied Economics Letters*, 12, 957-962.

[17] Kim, C.-J. and Nelson, C. R. (1999) *State-space Models with Regime Switching*, MIT Press: Cambridge.

[18] Krolzig, H. M. (1997), Statistical Analysis of Cointegrated VAR Processes with Markovian Shifts, manuscript.

[19] MacKinnon, J. G. (1991), Critical Values for Cointegration Tests, in: R. F. Engle and C. W. J. Granger (eds.) (1991), *Long-run Economic Relationships,* Oxford University Press, Oxford, 267-276.

[20] Martin, G. M. (2000), US Deficit Sustainability: a New Approach Based on Multiple Endogenous Breaks, *Journal of Applied Econometrics*, 15, 83-105.

[21] Payne, J. E. (1997), International evidence on the sustainability of budget deficits, *Applied Economics Letters*, 4, 775-779.

[22] Payne, J. E., Mohammadi, H. and Cak, M. (2008), Turkish budget deficit sustainability and the revenue-expenditure nexus, *Applied Economics*, 40, 823-830.

[23] Psaradakis, Z., Sola, M. and Spagnolo, F. (2004), On Markov Error-Correction Models, with and Application to Stock Prices and Dividends, *Journal of Applied Econometrics*, 19, 69-88.

[24] Quintos, C. E. (1995), Sustainability of the deficit process with structural shifts, *Journal of Business and Economic Statistics*, 13, 409-417.

[25] Stock, J. H. and Watson, M. W. (1993), A Simple Estimator of Cointegrating Vectors in Higher Order Integrated Systems, *Econometrica*, 61, 783-820.

[26] Wilcox D. (1989), The Sustainability of Government Deficit: Implications of the Present-value Borrowing Constraint, *Journal of Money Credit and Banking*, 21 291-306.

7 Appendix

Table 1: Cointegration analysis

Countries	AEG	PO	$\hat{b}$	AEG_{DOLS}	PO_{OLS}	$\hat{b}_{OLS}$
Bahamas	−3.460*	−12.575**	0.899 (0.042)	−3.324	−9.395**	0.972 (0.036)
Finland	−2.015	−3.555*	0.813 (0.042)	−1.609	−2.866**	0.824 (0.046)
France	−7.671**	−8.229**	0.825 (0.067)	−4.353**	−6.956**	1.013 (0.059)
South Africa	−3.670*	−7.049**	0.929 (0.035)	−2.836	−3.761*	0.975 (0.031)
Thailand	−1.725	−3.378*	1.005 (0.034)	−1.672	−2.519	1.008 (0.033)
United States	−2.469	−2.736**	0.969 (0.026)	−2.499	−2.196	0.978 (0.026)

Note: * significant at 5%, ** significant at 1%; standard errors in brackets

Table 2: Gregory-Hansen tests

	C/S			S			S_{nc}		
Countries	AEG	Z_α	Z_t	AEG	Z_α	Z_t	AEG	Z_α	Z_t
Bahamas	−3.95 [98:1]	−148.2* [92:3]	−13.46* [93:1]	−3.99 [98:1]	−133.8* [92:3]	−12.32* [92:3]	−4.01 [98:1]	−133.9* [90:4]	−12.32* [93:1]
Finland	−2.93 [99:2]	−30.2 [99:3]	−4.48* [99:3]	−3.71 [99:2]	−46.6* [98:1]	−5.67* [98:1]	−3.56 [96:2]	−46.9* [96:3]	−5.67* [96:3]
France	−9.26* [98:1]	−100.4* [98:1]	−9.30* [98:1]	−9.04* [98:1]	−97.5* [98:1]	−9.07* [98:1]	−9.42* [98:1]	−102.3* [98:1]	−9.46* [98:1]
South Africa	−4.40* [98:1]	−73.9* [98:1]	−7.67* [98:1]	−4.27 [98:1]	−67.3* [98:1]	−7.31* [98:1]	−4.40 [99:1]	−66.4* [98:1]	−7.26* [98:1]
Thailand	−2.16 [97:2]	−25.2 [97:1]	−3.79 [97:1]	−3.80 [95:3]	−54.6* [96:2]	−6.16* [96:1]	−4.60 [97:2]	−50.5* [96:2]	−5.85* [96:1]
United States	−3.96 [95:1]	−98.9* [95:2]	−8.91* [95:2]	−4.08 [95:1]	−102.4* [95:2]	−9.33* [94:4]	−4.78* [96:1]	−129.7* [97:2]	−12.01* [97:2]

*: rejection of the null hypothesis of no cointegration; potential break dates in square brackets

Table 3: Markov switching cointegration tests

Countries	AEG	PO
Bahamas	−4.608*	−12.555*
Finland	−8.593*	−8.879*
France	−7.922*	−8.227*
South Africa	−4.716*	−8.445*
Thailand	−4.709*	−4.900*
United States	−8.593*	−8.463*

*: rejection of the null hypothesis of no cointegration

Table 4: Markov switching cointegration estimates

Countries	α_1	α_2	β_1	β_2	ω_1	ω_2	p	q	AIC$^{(-)}$	SIC$^{(-)}$
Bahamas	-0.055 (0.069)	1.409 (0.187)	0.979 (0.040)	-0.889 (0.108)	0.152 (0.020)	-0.025 (0.043)	0.934 (0.049)	0.978 (0.024)	-29.82	-16.01
Finland	-0.858 (1.291)	3.925 (2.495)	1.036 (0.025)	-0.311 (0.038)	4.667 (0.440)	0.972 (0.857)	0.912 (0.047)	0.979 (0.013)	-109.9	-96.04
France	0.068 (0.027)	0.899 (0.115)	0.964 (0.046)	-1.309 (0.183)	0.039 (0.003)	-0.005 (0.005)	0.921 (0.046)	0.976 (0.012)	-48.91	-35.54
South Africa	14.74 (78.01)	33.57 (83.26)	0.901 (0.133)	-0.196 (0.154)	31.57 (3.492)	-2.757 (4.621)	0.976 (0.020)	0.955 (0.028)	-40.35	-26.54
Thailand	660.7 (122.9)	-711.4 (123.9)	0.891 (0.072)	-0.200 (0.077)	238.8 (25.23)	-193.8 (31.41)	0.973 (0.020)	0.975 (0.011)	-159.4	-145.6
United States	0.135 (0.182)	0.081 (0.190)	0.744 (0.249)	-0.205 (0.264)	0.089 (0.024)	-0.062 (0.023)	0.961 (0.023)	0.921 (0.034)	-64.11	-50.26

Standard errors in brackets; $(-)$: difference between Markov switching and linear model

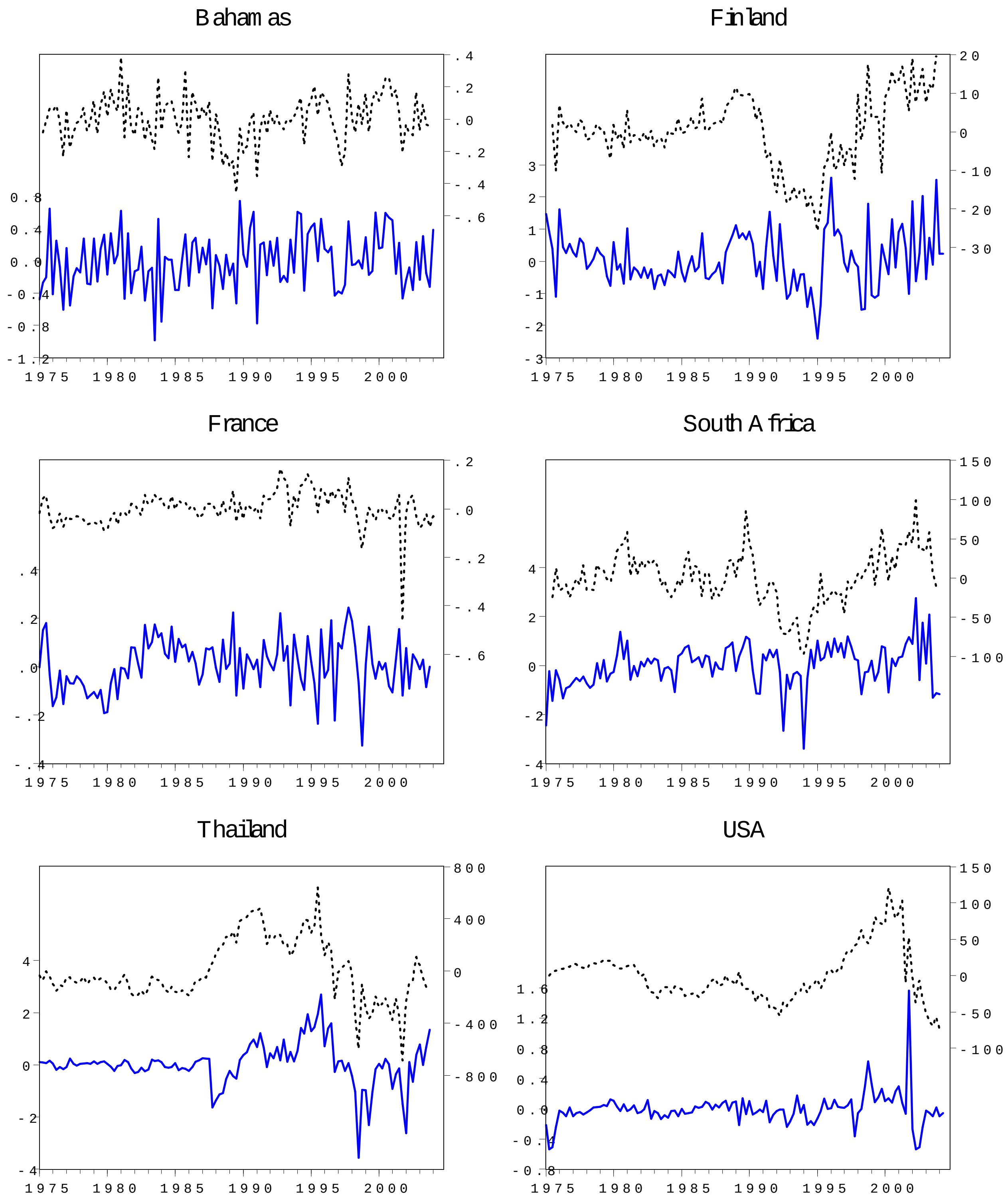

Figure 1: DOLS residuals (above) and standardized MS residuals (below)

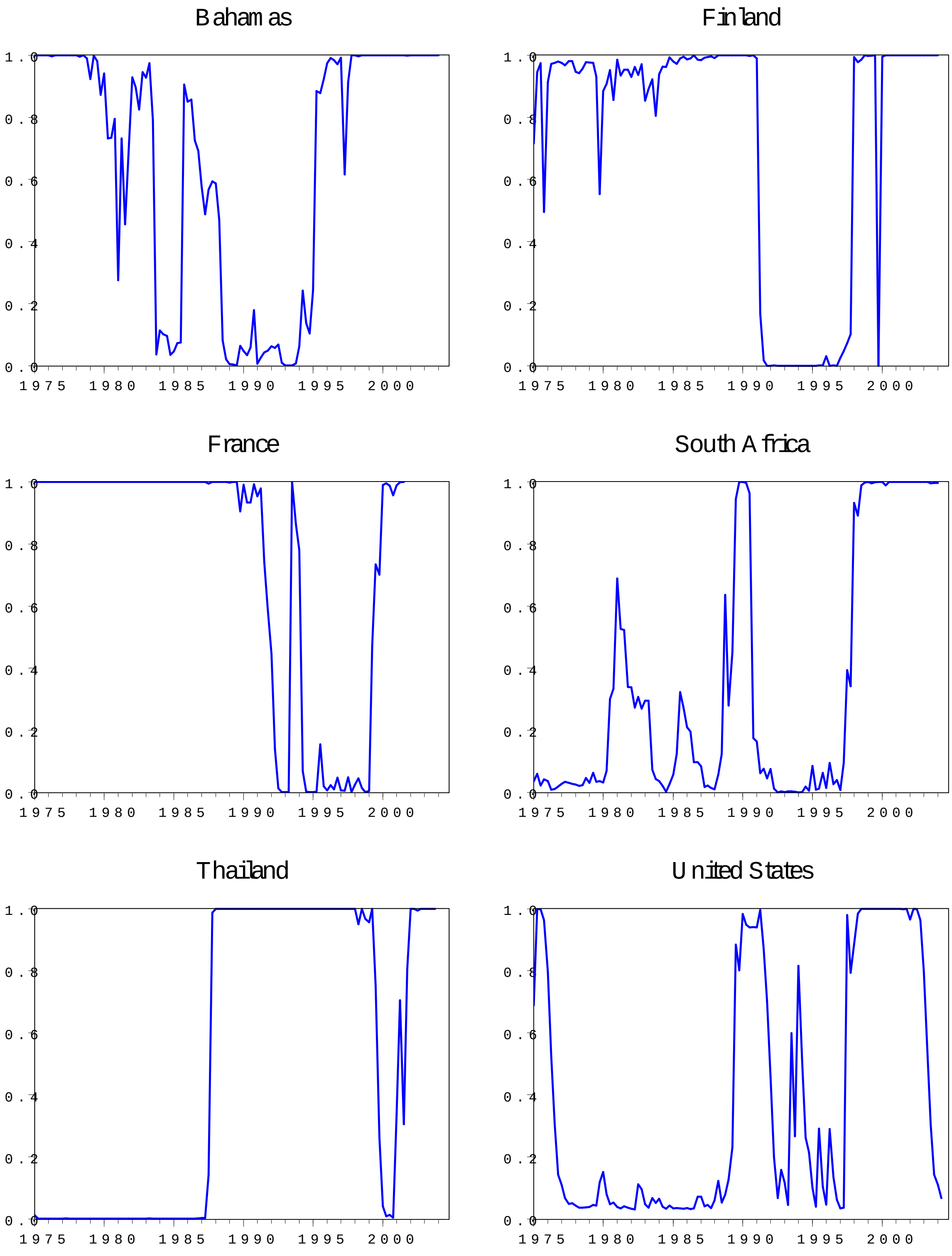

Figure 2: Regime 1 filtered probabilities

Most Recent Working Papers

NIPE WP 9/2009	**Gabriel,Vasco J. e Pataaree Sangduan,** "Assessing Fiscal Sustainability Subject to Policy Changes: A Markov Switching Cointegration Approach", 2009
NIPE WP 8/2009	**Aguiar-Conraria, Luís e Maria Joana Soares,** "Business Cycle Synchronization Across the Euro-Area: a Wavelet Analysis", 2009
NIPE WP 7/2009	**Afonso, António, Luca Agnello, Davide Furceri e Ricardo M. Sousa** "Assessing Long-Term Fiscal Developments: a New Approach", 2009
NIPE WP 6/2009	**Furceri, Davide e Ricardo M. Sousa,** "The Impact of Government Spending on the Private Sector: Crowding-out versus Crowding-in Effects", 2009
NIPE WP 5/2009	**Esteves, Rosa Branca,** "A Survey on the Economics of Behaviour-Based Price Discrimination", 2009
NIPE WP 4/2009	**Peltonen, Tuomas A., Ricardo M. Sousa e Isabel S. Vansteenkiste,** "Wealth Effects in Emerging Market Economies", 2009
NIPE WP 3/2009	**Afonso, António e Ricardo M. Sousa,** "The Macroeconomic Effects of Fiscal Policy in Portugal: a Bayesian SVAR Analysis", 2009
NIPE WP 2/2009	**Bastos, Paulo e Natália P. Monteiro**, "Managers and wage policies", 2009
NIPE WP 1/2009	**Afonso, António e Ricardo M. Sousa**, " Assessing Long-Term Fiscal Developments: Evidence from Portugal ", 2009
NIPE WP 22/2008	**Afonso, António e Ricardo M. Sousa**, "The Macroeconomic Effects of Fiscal Policy", 2008
NIPE WP 21/2008	**Afonso, António e Ricardo M. Sousa**, "Fiscal Policy, Housing and Stock Prices", 2008
NIPE WP 20/2008	**Magalhães, Pedro C., Luís Aguiar-Conraria,** "Growth, Centrism and Semi-Presidentialism: Forecasting the Portuguese General Elections", 2008
NIPE WP 19/2008	**Castro, Vítor,** "Are Central Banks following a linear or nonlinear (augmented) Taylor rule?", 2008
NIPE WP 18/2008	**Castro, Vítor,** "The duration of economic expansions and recessions: More than duration dependence", 2008
NIPE WP 17/2008	**Holmås, Tor Helge, Egil Kjerstad, Hilde Lurås and Odd Rune Straume,** "Does monetary punishment crowd out pro-social motivation? The case of hospital bed-blocking", 2008.
NIPE WP 16/2008	**Alexandre, Fernando, Miguel Portela and Carla Sá,** "Admission conditions and graduates' employability", 2008.
NIPE WP 15/2008	**Alexandre, Fernando, Pedro Bação and Vasco Gabriel,** "Taylor-type rules versus optimal policy in a Markov-switching economy", 2008.
NIPE WP 14/2008	**Bastos, Paulo, Natália P. Monteiro and Odd Rune Straume**, "Firm heterogeneity and wages in unionised labour markets: Theory and evidence", 2008.
NIPE WP 13/2008	**Esteves, Rosa Branca**, "Price Discrimination with Private and Imperfect Information", 2008.
NIPE WP 12/2008	**Esteves, Rosa Branca,** "Price Discrimination with Partial Information: Does it pay off?", 2008.
NIPE WP 11/2008	**Ribeiro, José Cadima and José Freitas Santos,** "Portuguese quality wine and the region-of-origin effect: consumers' and retailers' perceptions", 2008.
NIPE WP 10/2008	**Lommerud, Kjell Erik, Frode Meland and Odd Rune Straume,** "Mergers and capital flight in unionised oligopolies: Is there scope for a 'national champion' policy?", 2008.
NIPE WP 9/2008	**Gabriel, Vasco J., Paul Levine and Christopher Spencer,** "How forward-looking is the Fed? Direct estimates from a 'Calvo-type' rule", 2008.
NIPE WP 8/2008	**Brekke, Kurt R., Roberto Cellini, Luigi Siciliani and Odd Rune Straume,** "Competition and quality in regulated markets: a differential-game approach", 2008.
NIPE WP 7/2008	**Neto, Delfim Gomes, and Francisco José Veiga**, "Financial globalization, convergence and growth", 2008.
NIPE WP 6/2008	**Brekke, Kurt R., Luigi Siciliani and Odd Rune Straume,** "Hospital competition and quality with regulated prices", 2008.