The Exogeneity (at best) of the Optimum Currency Area Criteria for the Euro Zone*

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Abstract

A number of papers over the last decade have posited that Optimal Currency Areas are endogenous with respect to business cycle synchronization. The claim is that a common currency will greatly increase trade, and then trade will increase output synchronization. Countries that thus seem ill-suited for a common currency prior to a monetary union may become well-suited once the union is in place. There are other channels, however, through which a common currency could lead to convergent-or divergent- business cycles that this literature overlooks. We thus test directly for the impact of the Euro on business cycle synchronization using two variants of Markov Switching models. Examining four small Euro zone countries, our results indicate there was no increase in convergence with the larger economies of Germany and France in two cases, and in the other two cases, actually *divergence.* In contrast, when investigating three non-Euro zone countries, there is no significant loss of synchronization in two cases, and an increase in synchronization in one case. These results are precisely the opposite of what would be expected from the endogenous OCA literature, and bolster the case for caution before nations with differing cycles or structures enter a currency union.

JEL Codes: F33, F41

*We would like to thank participants at the 2011 Chinese Economic Association in North America Conference, 2011 Southern Economics Association Conference, 2011 Missouri Valley Economics Association Conference, University of Illinois Development/International Seminar, Richard Anderson, Nathan Balke, Werner Baer, Jen-Chi Cheng, Dong Cho and Hadi Esfahani for helpful comments.

Introduction

Preceding its beginning in 1999, the Euro zone has attracted skeptics who worried that a single monetary policy would be problematic for such a varied group of nations. For instance, in the Euro's early years there were "peripheral" nations such as Ireland and Spain, which grew fast, and experienced some asset and housing bubbles (which of course later burst), and would presumably have benefitted from tighter monetary policy over 1999-2006. In contrast, wealthier "center" nations such as France and Germany experienced very sluggish growth in the initial years of the currency and may well have benefitted from looser money and lower interest rates than the ECB provided. More recently, of course, many peripheral countries face major adjustment problems, with burst housing bubbles, large government and private sector indebtedness and very high unemployment, and loose monetary policy at the moment would be the standard prescription. Yet at the same time, the ECB has been criticized for being too tight in the face of the current global downturn (Angela Merkel, Germany's prime minister, has expressed a desire that the ECB run an even *tighter* policy, and two ECB officials from Germany-Juergen Stark and Axel Weber-have resigned in apparent protest of central bank attempts to alleviate the debt crisis by buying sovereign bonds). The Euro zone, in short, may not have been an Optimal Currency Area (OCA).

Traditional criteria for joining a currency union include factor mobility between the prospective currency union members (see Mundell, 1961), as well as the extent of trade. Later economists added such standards as a system of risk sharing through fiscal policy, and, especially for the purposes of a "one-size-fits-all" monetary policy, highly synchronized business cycles (see Alesina and Barro, 2002). The lack of synchronized business cycles between the center and peripheral countries of the Euro zone may be the most important reason for doubts that the now seventeen member union is truly an OCA.

On the other hand, a growing academic literature over the last decade (see Rose, 2008 for a summary and meta-analysis) has produced results indicating that, even if a currency union such as the Euro-zone isn't an optimal currency area prior to nations joining it, the very act of creating and joining the common currency makes the union an OCA ex-post. There are two mechanisms at work. First, joining a common currency will increase trade-the initial estimates of Rose suggest a huge effect; currency unions appeared to triple trade. Secondly, the increase in trade will increase business cycle synchronization among the different nations in the currency union. If these two mechanisms work, then even if different countries seem ill-suited for currency union ex-ante, they will, upon joining the union, experience more synchronized business cycles, making one currency (and one central bank and monetary policy) optimal. In this way, an OCA is endogenous.

While the results of Rose and others are provocative, there are a number of empirical problems with their methodology. First, the methods for estimating the effects of currency unions on trade, and trade on synchronization have been subject to much criticism. Secondly, Rose (2008), in his meta-analysis, combines estimates of the impact of currency unions on trade with estimates of the effect of trade on business cycle synchronization to get an estimate of the total impact of joining an OCA on synchronization. This technique assumes that trade is the only channel through which a currency union affects synchronization. To cite one counter-example, if a currency union increases capital flows, it could send investment from stagnant, center countries to peripheral nations, helping to fuel bubbles in the latter, thus leading to less synchronization. There could be several important ways in which a common currency could affect business cycle synchronization. The goal of this paper is not to list or develop models on all such channels. Rather, we will employ a technique which will directly estimate the impact of "center" output on "peripheral" business cycles.

The purpose of this paper is to use the nonlinear Markov Switching (MS) technique to check if four peripheral Euro zone nations have become more synchronized with the center nations in their output growth. [1] We will then, as a further check, examine whether three smaller European countries, which have chosen not to adopt the Euro, have experienced a decrease in business cycle synchronization with

[1] This Markov-Switching (MS) technique has been employed elsewhere to model business cycles and examine the impact of cyclical shocks between countries (Santos, 2002).

the center of Euro zone as a result of retaining their own currencies. Both the time-varying transitional probability (TVTP) and the traditional fixed transitional probability (FTP) MS models will be used to examine the output dynamics of these small economies. The TVTP model assumes that certain economic variables, such as the center nations' output growth, affect the transitional probability. In contrast, the FTP model assumes that the transitional probabilities are exogenous and hence the probabilities are not modeled as functions of economic variables. In an FTP model, the unobserved state variable, dictated by constant probabilities, determines the state of the economy, i.e., being in an expansion or recession. Based on the FTP estimation results, we are able to derive the probability of the economy being in recession or expansion. Instead of incorporating variables such as the center nations' output growth directly into the original FTP equation, we use a probit model to investigate the impacts of the center nations' output growth on the peripheral nation's probability of being in a recession.[2] Though these TVTP and FTP models are different, they provide a consistent insight about this topic.

Our results cast serious doubt on hopes that the smaller nations of the Euro-zone have made progress toward being part of a European OCA by the mere act of joining. Results differ slightly among particular countries between the two different Markov Switching models employed, but are consistent in two respects. First, for both MS models, in two of the four smaller Euro-zone economies, there is no significant increase in synchronization since adopting the Euro. More importantly, for the other two Euro zone nations, there has been a significant *decrease* in synchronization with the large European economies since entering the Euro-zone.

Secondly, for the three non-Euro zone nations, in two cases there was no significant change in synchronization with the larger economies of Europe. For the third nation -Sweden in the TVTP MS

[2] As will be shown later, we use five explanatory variables in the probit model. Should the five variables be included in the original FTP model, it creates difficulty in obtaining convergent estimates for some countries. In some cases, even if the convergence is achieved, the Hessian matrix fails to be positive definite at the point of supposed convergence, probably due to overflow and/or underflow problems and the ensuing numerical inaccuracies. To circumvent this barrier, we use FTP/probit approach. In our FTP/probit setting, we assume that variables such as center nations' output growth rates are contained in the error term. Since the error term drives the economy to be in recession or in expansion, we then examine how the second stage variables, such as the center nations' output growth rates, affect these phenomena.

specification and Denmark in the FTP -there has been a significant increase in synchronization with central Europe since refusing to join the Euro zone. Thus in these cases failure to adopt the Euro has not been sufficient to reverse integration with the main economies of the continent, while for two Euro zone countries joining the Euro has failed to prevent a divergence in cyclical fluctuations with its currency partners.

This paper proceeds as follows. The next section details the literature on OCAs and their possible endogeneity. The third section discusses the methodology, and the fourth presents our results. The fifth section concludes.

Previous Literature

The classic work on OCA goes back to Mundell (1961) with contributions by among others, McKinnon (1963) and Kenen (1969). The early literature emphasized the importance of factor mobility and the level of trade as criteria for whether a group of countries should form a currency union. Later work added a system of fiscal transfers and the similarity of shocks and business cycles as conditions for a common currency to be truly "optimal". These last two metrics could be vital-once a currency union (CU) is in place, there is only one monetary policy. Thus it could be problematic if potential members of a CU exhibited asymmetric responses to shocks and little output co-movement.

Given the different sizes and structures of the varied Euro-zone economies, business cycle synchronization was a major concern prior to the Euro's existence. However, a criticism of business cycle synchronization, as well as other criteria for OCAs, is that they are potentially endogenous. For instance, it may well be the case that the extent of trade between countries will be different before and after adoption of a CU. Starting in 1996, a series of papers beginning with Frankel and Rose (1996) presented results indicating that joining a CU greatly increases trade. Indeed, the initial estimates of the impact of a common currency on trade suggested a CU would triple trade between nations sharing the same money.

Moreover, the authors presented results indicating that increased trade raises business cycle synchronization. In principle, standard Ricardian or Hecksher-Ohlin models would suggest that greater trade would lead to greater specialization, and thus nations would exhibit more asymmetric responses to shocks. However, as Frankel and Rose (1998) point out, "if demand shocks (or other common shocks) predominate, or if intra-industry trade accounts for most trade, then business cycles may become *more* similar across countries when countries trade more. We believe this latter case to be the more realistic one.." (p. 1010). The authors indeed present results indicating that trade raises business cycle synchronization (BCS).

These results on the endogeneity of the OCA criteria seem to bolster the case for common currencies, even between nations at different stages of development. First, joining the CU raises trade, and, secondly, the tighter trade links increase BCS. In principle, countries not ready *ex-ante* for a CU can be made ready *ex-post* by the very act of joining the CU. To quote from a 1996 paper regarding the case of Sweden and its possible membership in the Euro zone, "Trade patterns and income correlations are endogenous. Sweden could fail the OCA criterion for membership today, and yet, if it goes ahead and joins anyway, could, as a result of joining, pass the Optimum Currency Area (OCA) criterion in the future" (Frankel and Rose, 1996, p. 1).

There are, however, several problems of interpretation. First, the impact of a CU on trade, while estimated to be very large by Rose, has been found to be much smaller by other researchers. Persson (2001), for instance, posits that countries which join CUs do not do so randomly; rather, there are factors which drive countries both to trade and join CUs. By applying a matching estimator to the data, the author finds that the effect of a CU on trade is statistically insignificant. Similarly, Wolf and Ritschl (2011) examine the trade effects of common currencies by investigating a number of post-Depression currency blocs. Standard estimates of the impact of such currency blocs indicate large trade effects; but the authors go on to find the estimates are biased by endogeneity. Wolf and Ritschl conclude by cautioning against accepting the high "trade creation effects" of arrangements such as common currencies often found in the literature. The controversy over the impact of CUs on trade has persisted for over a decade.

Moreover, even if CUs were to increase trade and greater trade in itself raises BCS, the effect of greater trade on BCS may be swamped by other factors. For instance, if joining the common currency leads to a sharp increase in foreign borrowing (see Eichengreen and Hausman, 1999, pp. 91-92 for a discussion on exchange rate rigidity and foreign borrowing), then capital may flow from relatively stagnant "center" countries to faster growing "peripheral" nations with higher returns, fueling growth (and perhaps bubbles) in the peripheral nations. This process, which would describe the experience of several smaller Euro-zone countries in the early years of the currency, would tend to lower, rather than raise BCS (especially when the housing and asset bubbles in the periphery burst, and capital flows quickly reversed). Indeed, Alesina, Barro and Tenreyo (2002) find that CUs do raise trade, but do not, in most specifications, have a significant impact on output co-movement. Using a panel approach, and attempting to directly address the issue of endogeneity between trade and common currencies, Barro and Tenreyo (2007) find that trade is enhanced by CUs, while the co-movement of output shocks actually *decreases* in response to sharing the same currency. Along these lines, Buscher and Gabrisch (2011) find that the Euro has had little impact in synchronizing nominal wage dynamics in the Euro zone, as asymmetries in nominal wage formation continue to persist across the continent's common currency area.

Another way in which a CU could decrease, rather than increase BCS is through different inflation rates in member countries. With one policy interest rate, low inflation countries have a higher real interest rate than higher inflation nations. The Economist quoted Patrick Artus in 2005, a period of sluggish growth in Germany but fast growth in the periphery, on this phenomenon. "Because all member states share the same nominal interest rate, slow-growing economies with lower inflation, such as Germany and Italy, have higher real interest rates than fast growers, such as Spain and Greece. This is the exact opposite of what is needed, exacerbating the divergence in growth rates" (Economist, August 18th 2005, p. 18).

A further difficulty of interpretation with the Frankel and Rose results is that most studies of CU and BCS measure BCS as a correlation coefficient "that is estimated between detrended levels of activity for countries i and j" (Rose, 2008, p. 6). Even assuming the detrending techniques-such as Hodrick-Prescott or Baxter-King decompositions-are reliable, they yield one estimate for an entire cross-country sample. These are of course subject to the same criticisms of cross-country studies on other topics (see Eichengreen, 2001), and even if their average estimate is an unbiased statistic for the average effect, individual country experience may differ substantially from the average[3].

Despite the difficulties of interpretation, Rose (2008) attempts to gauge the impact of CUs on BCS by performing two meta-analyses-one for the effect of CUs on trade, and one for the impact of trade on BCS. The author combines the two results of the meta analyses, and states that the effect of CUs on BCS is "substantial, whether it is enough to obviate the need for a national monetary policy is of course a different question". Unfortunately, even here the finding of a "substantial" impact of CUs on BCS may be unwarranted. Both meta-analyses rely on cross-country studies, with all of their aforementioned attendant problems. Moreover, this method, by focusing only on trade and BCS, ignores other channels through which CUs can affect BCS.

Interestingly, Willett, Permpon and Wihlborg (2010) examine the Euro and BCS through the correlation of output and consumption, both within Euro zone countries and within non-Euro zone European nations. They find that the growth in the correlation of output was greater for European countries outside the Euro zone than for those within the currency union! This result is robust to measuring output shocks as deviations from a Hodrick-Prescott trend. There may of course be questions

[3] On a related note, Alesina, Barro and Tenreyro (2002) point out that if a CU is between a large (anchor) and small (client) nations, "What turns out to matter is not the correlation of shocks, per se, but rather the variance of the client country's output expressed as a ratio to the anchor country's output. This variance depends partly on the correlation of output (and, hence, of underlying shocks) and partly on the individual variances of outputs. For example, a small country's output may be highly correlated with that in the United States. But, if the small country's variance of output is much greater than that of the United States, then the U.S. monetary policy will still be inappropriate for the client. In particular, the magnitude of countercyclical monetary policy chosen by the United States will be too small from the client's perspective." (pp. 7-8).

of interpretation with using output correlations to measure BCS, as noted by Alesina, Barro and Tenreyo (2002) (see footnote 3). But the Willett, et al. findings remain striking.

Given all of the problems with the cross country methods, and the very interesting findings of Willett, et al., we thus propose a different method for examining directly whether the Euro has led to greater BCS in peripheral countries. Unlike cross-country studies which estimate one effect for all nations, our method allows a different impact of the Euro on the synchronization of different smaller nations with the larger economies of the Euro-zone.

Data and Methodology

Our goal is to determine how much, if at all, the Euro has increased BCS between the larger and smaller economies of this most famous modern currency area. If Rose, et al. are correct, we should observe an increase in BCS between the large and small EMU countries, and no such increase (or at least a smaller increase) in BCS between center EMU nations and non-Euro zone European states. Failure to find such a pattern would be evidence against the proposition that the very act of joining the Euro increases the suitability of the common currency. In addition, we will use the United States to represent the rest of the world and examine how synchronized the peripheral EMU and non-EMU European countries have become with the business cycle of the United States during the Euro's existence. Again, if Rose, et al. are correct, we should observe the smaller EMU countries becoming more synchronized with Germany and France, with less relative synchronization with the US since Euro adoption.

The data sets used in this paper are quarterly seasonally adjusted industrial production (IP) from 1983:Q1 to 2009:Q4. Data were obtained from the International Financial Statistics database of the IMF. We take the log difference as our output growth measure. We define the center and the periphery as follows: the center countries are Germany and France. The peripheral Euro-zone countries are Greece, Ireland, Portugal and Spain.[4] The peripheral non-EMU countries that we will examine for comparison are Denmark, Switzerland and Sweden.

[4] Even though most people are interested in the so-called "PIIGS", Italy is not included in this study due to the large size of its economy.

For our first exercise, we run simple linear regressions of peripheral country output growth with the following specification:

$$y_t = \delta_0 + \delta_1 D_{99} + \delta_2 y_{GF} + \delta_3 D_{99} y_{GF} + \delta_4 y_{US} + \delta_5 D_{99} y_{US} + \varepsilon_t$$

where y_t is the log change in peripheral country output; $D_{99} = 1$ if $t \geq$ 1999Q1 while $D_{99} = 0$ if t < 1999Q1; y_{GF} and y_{US} are log changes in German-French and US output, respectively. If the estimate of δ_3 is positive and significant, center output has increased its impact on the periphery since Euro adoption. This would be evidence in favor of the Euro being an endogenous OCA. An insignificant estimate of δ_3 would indicate no impact of Euro adoption on BCS. Since the peripheral country's output growth may also be affected by the rest of the world, we also include the United States output growth, and its interaction with the Euro years in this equation.

It is important to test for nonlinearity. Though there are many procedures to test against certain particular nonlinear models such as the Threshold Autoregressive or Smooth Transition Autoregressive specifications, we use the Brock, Dechert, and Scheinkman (BDS) test to examine non-linearity. BDS is a popular test because of its generality and usefulness against a variety of nonlinear time series models. The idea behind this test is that, if a given series is indeed linear, the residuals from the linear model should be i.i.d. That is, the probability that the distance between any two residuals is less than a given constant (denoted as epsilon) should be the same for all residuals. The rejection of the null implies that the series may be non-linear. We performed this test, using epsilons of 0.5, 1, 1.5, and 2 standard deviations of each data set; we find that the null hypothesis of the data being iid is rejected for Portugal, Spain, Denmark, Switzerland, and Sweden. For Greece, if the epsilons are 1.5 and 2 standard deviations of the data set, the null is rejected. For Ireland, the null cannot be rejected. When considering the BDS result for Ireland, it is important to keep in mind that this particular test has been shown to lack good finite sample properties (see Enders, 2010, p. 437).

Given the results of the nonlinearity tests, the exercise in simple regressions will yield evidence that is at best suggestive, as we are merely looking at linear estimates of the business cycle, which is an inherently nonlinear process. To gather evidence in a more rigorous fashion, we will employ Markov Switching (MS) models, which since Hamilton's (1989) pioneering paper, have been utilized to investigate business cycle fluctuations.

We construct our MS model for country output by starting with an AR(p) model,

$$y_t = \alpha_0 + \beta_1 y_{t-1} + \beta_2 y_{t-2} + \beta_3 y_{t-3} + ... + \beta_p y_{t-p} + u_t$$

where y_{t-i} is output growth at time $t-i$ for i=0, 1,2,.. We then transform it by arranging the terms and get the error correction format,

$$\Delta y_t = \alpha_0 + (\rho - 1) y_{t-1} + \gamma_1 \Delta y_{t-1} + \gamma_2 \Delta y_{t-2} + ... + \gamma_{p-1} \Delta y_{t-p+1} + u_t$$

where $\rho = \sum_{i=1}^{p} \beta_i$ and ρ is a measure of persistence.[5]

To evaluate the asymmetric properties of output dynamics during economic expansions and contractions, we introduce a state variable S_t, which denotes the state of the economy at time t. There are two distinct states: regime 0 (i.e., S_t=0) and regime 1 (S_t=1). The MS model can simultaneously handle changes in the mean as well as the variance of output growth[6]. The previous model can thus be expressed as

$$\Delta y_t = (\alpha_0 + \alpha_1 S_t) + \phi_0 y_{t-1} + \gamma_1 \Delta y_{t-1} + \gamma_2 \Delta y_{t-2} + ... + \gamma_{p-1} \Delta y_{t-p+1} + u_{t,S_t}$$

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[5] We will use an AR(3) model to demonstrate the equality of the following equations. For p > 3, the procedure is the same. Note that the equation $y_t = \alpha_0 + \beta_1 y_{t-1} + \beta_2 y_{t-2} + \beta_3 y_{t-3} + u_t$ is exactly the same as $\Delta y_t = \alpha_0 + (\rho - 1) y_{t-1} + \gamma_1 \Delta y_{t-1} + \gamma_2 \Delta y_{t-2} + u_t$ where $\rho = \sum_{i=1}^{3} \beta_i$, $\gamma_1 = -\beta_2 - \beta_3$, and $\gamma_2 = -\beta_3$. Note that in a traditional AR model, $\sum_{i=1}^{3} \beta_i$ is a measure of persistence.

[6] Allowing for changes in persistence, as well as the mean and variance, is possible. However, in doing so, the program has difficulty in achieving convergence for quite a few countries.

Where $\phi = \rho - 1$. A two-state Markov process has the following transition probabilities: $\Pr(S_t = 0|S_{t-1} = 0, \Phi_{t-1}) = \widetilde{q_t}$ and $\Pr(S_t = 1|S_{t-1} = 1, \Phi_{t-1}) = \widetilde{p_t}$ for the time-varying probability case; $\Pr(S_t = 0|S_{t-1} = 0) = \tilde{q}$ and $\Pr(S_t = 1|S_{t-1} = 1) = \tilde{p}$ for the constant probability case.[7] The parameter α_1 captures the change in the mean of output growth during regime 1 relative to regime 0. If α_1 is negative, regime 1 is the contraction regime while regime 0 is the expansion regime. Otherwise, regime 1 is the expansion regime and regime 0 is the contraction regime. For the variance, $Var(u_t) = h_0$ when $S_t = 0$ and $Var(u_t) = h_1$ when $S_t = 1$. In this paper, we will use $p = 3$.[8]

As mentioned, there are two ways to deal with the transitional probability. One is the TVTP model which assumes that the probabilities are affected by economic variables. The other is the FTP model which assumes constant probabilities. For the TVTP model, it would be ideal to incorporate all the following variables ($D_{99}, y_{GF,t-1}, D_{99}y_{GF,t-1}, y_{US,t-1}, D_{99}y_{US,t-1}$) as predictors affecting the transitional probabilities. However, the TVTP model has great difficulty in providing meaningful results for any country if there are more than two explanatory variables. This should not be surprising. Filardo (1994) pioneered the use of the TVTP MS model, but never used more than two variables (three if additional lags are counted) as determinants of the transition probabilities, and obtained mixed results in terms of significance. Santos (2002) faced similar challenges when estimating a TVTP model for Mexican output. The author employed numerous variables to explain the Mexican business cycle, but each regressor was entered into the TVTP MS model one at a time, rather than simultaneously. Moreover, despite trying

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[7] Since the latter is a more general case than the former, from now on, the mention of the latter implicitly includes the former.

[8] Due to the programming complexity in dealing with various state variable statuses at different time points, it is not trivial to have a model with a long lag length. As the lag length of the model increases, the time-dimension of the state variable increases substantially. For example, if the lag length is l (= p - 1 as shown in the equation above), the time-dimension of the state variable is a $2^{l+1} \times l+1$ matrix. Thus, for a lag length of $l = 2$, the time-dimension of the state variable G is an 8×3 matrix, i.e., $G' = \begin{bmatrix} 0&0&0&0&1&1&1&1 \\ 0&0&1&1&0&0&1&1 \\ 0&1&0&1&0&1&0&1 \end{bmatrix}$ where the three columns of G represent the statuses of the state variable at t-2, t-1, and t respectively. Then for $l = 3$, G is a 16×4 matrix; for $l = 4$, G is a 32×5 matrix.

each variable one at a time, the author found only one of the candidate variables showed significance in the TVTP model.

Given the difficulties in obtaining meaningful results for some countries[9] when we include all five predictor variables in the TVTP model, we will follow two approaches. First, to obtain convergent estimates of a TVTP MS model, we will drop the US variables ($y_{US,t-1}$ and $D_{99}y_{US,t-1}y_{US}$) and the Euro dummy (D_{99}) as we cannot obtain convergence when they are included in the model. We will then examine the impact, in the TVTP estimation, of $y_{GF,t-1}$ and $D_{99}y_{GF,t-1}$. Our second approach will be to estimate an FTP MS model, and then to run a probit model in which the probability of being in regime 1 is estimated as a function of the five variables above, for those countries in which the two US variables appear significant.

For the TVTP model, we will thus specify the transition probabilities as follows:

$$\widetilde{p_t} = \frac{\exp(\theta_{\mathrm{p}0} + \theta_{\mathrm{p}1}y_{GF,t-1} + \theta_{\mathrm{p}2}D_{99}y_{GF,t-1})}{1 + \exp(\theta_{\mathrm{p}0} + \theta_{\mathrm{p}1}y_{GF,t-1} + \theta_{\mathrm{p}2}D_{99}y_{GF,t-1})}$$

$$\widetilde{q_t} = \frac{\exp(\theta_{\mathrm{q}0} + \theta_{\mathrm{q}1}y_{GF,t-1} + \theta_{\mathrm{q}2}D_{99}y_{GF,t-1})}{1 + \exp(\theta_{\mathrm{q}0} + \theta_{\mathrm{q}1}y_{GF,t-1} + \theta_{\mathrm{q}2}D_{99}y_{GF,t-1})}$$

Thus for the TVTP, the parameters are $(\alpha_0, \alpha_1, \phi_0, \gamma_1, \gamma_2, h_0, h_1, \theta_{\mathrm{p}0}, \theta_{\mathrm{p}1}, \theta_{\mathrm{p}2}, \theta_{\mathrm{q}0}, \theta_{\mathrm{q}1}, \theta_{\mathrm{q}2})$. For our fixed transition probability approach, we note that the parameters of the FTP model are $(\alpha_0, \alpha_1, \phi_0, \gamma_1, \gamma_2, h_0, h_1, \tilde{p}, \tilde{q})$. MLE is implemented to estimate the parameters. The log-likelihood function is the following:

$$\ln L = \sum_{t=1}^{T} \ln \left\{ \sum_{S_t=0}^{1} \sum_{S_{t-1}=0}^{1} \sum_{S_{t-2}=0}^{1} f(\Delta y_t \mid S_t, S_{t-1}, S_{t-2}, \Phi_{t-1}) \cdot \Pr(S_t, S_{t-1}, S_{t-2} \mid \Phi_{t-1}) \right\}$$

[9] Convergence is difficult to achieve. As mentioned earlier, in some cases, convergence is achieved but the Hessian matrix fails to be positive definite. Thus, we could not obtain the standard errors of some estimates.

where

$$f(\Delta y_t|S_t,S_{t-1},S_{t-2},\Phi_{t-1}) = \frac{1}{\sqrt{2\pi\sigma_{s_t}^2}} exp\left\{-\frac{(\Delta y_t-(\alpha_0+\alpha_1 S_t)-\phi_0 y_{t-1}-\gamma_1\Delta y_{t-1}-\gamma_2\Delta y_{t-2})^2}{2\sigma_{s_t}^2}\right\}$$

and

$$\Pr(S_t = j, S_{t-1}=i, S_{t-2}=k \mid \Phi_{t-1}) = \Pr(S_t = j \mid S_{t-1}=i, S_{t-2}=k, \Phi_{t-1})\cdot\Pr(S_{t-1}=i, S_{t-2}=k \mid \Phi_{t-1})$$

where Φ_{t-1} denotes past information, $\sigma_{S_t}^2 = h_0(1-S_t)+h_1 S_t$ and the transition probability, for *i*, *j*, *k*=0,1, is $\Pr(S_t = j|S_{t-1}=i, S_{t-2}=k,\Phi_{t-1}) = \Pr(S_t=j|S_{t-1}=i,\Phi_{t-1})$.

A special characteristic of the FTP MS model is that we are able to calculate the probability of the economy in S_t=1 for each time period. Based on these probability values, we will assess the impacts of German-French output growth on the business cycles of peripheral Euro and non-Euro members, both before and after the creation of the Euro. We will implement the following Probit model for each of the seven small nations:

$$F^{-1}\big(P(S_t=1)\big) = \tau_0+\tau_1 D_{99}+\tau_2 y_{GF}+\tau_3 D_{99}\cdot y_{GF}+\tau_4 y_{US}+\tau_5 D_{99}\cdot y_{US}+\epsilon_t$$

As noted, y_{GF} and y_{US} are German-French and U.S. industrial production growth rates, respectively, and D_{99} measures the post-1998 Euro quarters. The coefficients τ_2 and τ_4 thus measure the effect of German-French and U.S. output growth, respectively, on the probability of recession[10] in the seven smaller Euro and non-Euro countries. The interaction coefficients τ_3 and τ_5 are the additional impacts of German-French and U.S. output growth on the probability of recession post-Euro adoption. If the theory of endogenous OCAs is correct, τ_3 should be negative and significant for the four smaller Euro economies. That is, since Euro adoption, an increase in German-French output should lower the

[10] Regime 1 is a recession regime when $\alpha_1$is negative. Later results indicate that most nations have negative estimates of α_1.

probability of recession in the smaller Euro economies more than before the creation of the common currency.

Results

Table 1 contains results for the simple regressions. In all four cases (Greece, Ireland, Portugal and Spain), the estimate of δ_3-interaction coefficient with the Euro era and German-French output growth- is insignificant. These results indicate that, based on simple linear regression, none of the four Euro zone countries in the sample has experienced an increase in BCS with the center of the monetary union. On the other hand, for the three non-Euro zone countries, the interaction term is insignificant in two of three nations (Denmark and Switzerland) but positive and significant (at the ten percent level) for Sweden. That is, despite failure to join the Euro zone, Sweden's economy appears more affected by French-German output changes since the Euro's creation. This is in contrast to all four Euro zone countries which display no such increase in BCS with the center of the Euro zone since adopting the common currency.

It may be questionable to include U.S. output growth in the model. We therefore test the restriction $H_0\colon \delta_4 = \delta_5 = 0$. For Portugal, Ireland and Spain, our results indicate that the inclusion of U.S. output growth is relevant, at least at the ten percent level. And indeed, while none of the Euro-zone nations have become more synchronized with the center of the CU, both Portugal and Ireland appear to have become less synchronized with the United States.

We also test for $H_0\colon \delta_2 = \delta_3 = 0$. As displayed in Table 1, at the 10% significance level, all except Ireland display a responsiveness to the Euro zone center economy. Finally, the test results of $H_0\colon \delta_2 = \delta_3 = \delta_4 = \delta_5 = 0$ are reported on the last row of Table 1, and, as displayed, the null is rejected for all countries. Of course the results of these regressions are really nothing more than informal linear exercises, so any inference must be very tentative. We thus turn to investigating how BCS has changed since the Euro by applying the MS models.

Table 2 shows the results of our TVTP model. For all countries, regime 1 is more volatile than regime 0. As for the average IP growth rates, all countries, except Greece, have negative estimates of α_1, i.e., regime 1 is the contraction regime for all countries except Greece. Thus, for Greece only, regime 1 is the expansion regime. Note that θ_{qi} where *i*=1,2,3 are related to regime 0 while $\theta_{pi}s$ are related to regime 1. In particular, θ_{p2} measures the impact of the interaction of French-German output changes and the Euro years on the probability of staying in recession for all countries except Greece (where it measures the impact of this interaction on the probability of staying in an expansion). Analogously, θ_{q2} measures the impact of this interaction term on the probability of staying in recovery for all countries, except for Greece (where it measures the impact on the probability of staying in a recession). None of the estimates of θ_{q2} is significant. For Euro zone members, the estimates of θ_{p2} are insignificant for Greece and Spain, which indicates no increase in BCS since joining the Euro zone. Notably, however, θ_{p2} is positive and significant for both Portugal and Ireland. That is, an increase in French-German output, since the Euro's introduction, raises the likelihood of a recession in these two Euro-zone nations-a result exactly the opposite of what would be expected if the OCA criteria were endogenous! Joining the Euro has apparently led to less, not more synchronization for these two smaller Euro-zone members.

For the non-Euro countries, θ_{p2} is insignificant for Denmark and Switzerland, indicating no loss in BCS from retaining their own currencies. However, the coefficient is negative and significant for Sweden, indicating an increase in BCS with the Euro-zone's center since the Euro's introduction, despite Sweden's failure to join the Euro zone. Thus, none of the four Euro zone nations in the sample has experienced an *increase* in BCS since the currency's introduction, and in two cases there has been a *decrease* in BCS. None of the three non-Euro zone nations has experienced a decrease in BCS since the Euro's introduction, and in one case there has been an *increase* in BCS. These results are precisely the opposite of those predicted by the endogenous OCA literature.

Even with these results, we want to test whether the TVTP model is preferred relative to the FTP specification. In the last row of Table 2, we impose the restriction of $\theta_{\mathrm{p}1} = \theta_{\mathrm{p}2} = \theta_{\mathrm{q}1} = \theta_{\mathrm{q}2}$=0. This is the restriction that transitional probabilities are not affected by any economic variables. By using the likelihood values of both the unrestricted and restricted cases,[11] the likelihood ratio test of $H_0{:}\, \theta_{\mathrm{p}1} = \theta_{\mathrm{p}2} = \theta_{\mathrm{q}1} = \theta_{\mathrm{q}2}$=0 indicates that the null hypothesis cannot be rejected for all countries, except Greece and Switzerland. Thus, compared to the FTP, TVTP is actually not a clearly preferred model for Portugal, Ireland, Spain, Denmark and Sweden. The failure to demonstrate the significant impact of these variables on the transitional probabilities in these countries does not, of course necessarily imply that these variables are irrelevant in these small economy's business cycles. Thus, for a more robust investigation, we will now turn to the FTP model.

Table 3 provides the estimates of the FTP MS model. For all nations, the estimate of α_0 is positive. The estimate of α_1 signals a shift in the mean of output growth as S_t varies from 0 to 1. Just as was the case for the TVTP model, the estimate of α_1 is negative for all countries except Greece. In other words, regime 1 is a contraction regime and regime 0 is the expansion regime for all countries except Greece. Thus, for the period of 1983Q1-2009Q4 in Ireland, Sweden and Switzerland, there is a significant decline in the mean industrial production growth rate in regime 1 (i.e., when S_t is one). The long-run average industrial growth rates in expansion for Portugal, Ireland, Greece, Spain, Denmark, Sweden and Swizerland are 2.52%, 10.87%, 0.569%, 2.097%, 3.190%, 2.660%, and 5.047% while the average growth rates in recession are –0.91%, 6.74%, 0.32%, 0.69%, 1.15%, -51.21%, and -0.51%.[12]

[11] In the case of TVTP, since the transitional probabilities depend on the lagged values of other economic variables, the calculation of the likelihood value of the model starts from the 4th observation (see Kim and Nelson 1999, p.93, Program 5). This differs from the FTP model where the calculation of the likelihood value could start earlier. Thus, the likelihood value of the FTP model is different from that of the TVTP model, even when $\theta_{\mathrm{p}1} = \theta_{\mathrm{p}2} = \theta_{\mathrm{q}1} = \theta_{\mathrm{q}2}$=0. Though it seems to be a programming issue, it actually becomes important in implementing the likelihood ratio test in the choice of TVTP versus FTP. A correct test of $\theta_{\mathrm{p}1} = \theta_{\mathrm{p}2} = \theta_{\mathrm{q}1} = \theta_{\mathrm{q}2}$=0 should use the same program with the same number of observations.

[12] The average rate is calculated as $(\alpha_0 + \alpha_1 S_t)/(1-\rho)$ where $\phi_0 = \rho - 1$.

In regard to the dynamics of output variability, the estimates of h_0 and h_1 indicate the volatility in regimes 0 and 1, respectively.[13] With the exception of Denmark and Sweden, both estimates are statistically significant. For all countries, the estimate of h_1 is greater than that of h_0. Thus, for all countries except Greece, there is a substantial increase in industrial production variability as the economy moves from the expansion regime (i.e., regime 0) to the recession regime (i.e., regime 1). This accords with Mitchell's (1927) claim that downturns tend to be more volatile than expansions. The exception here is Greece. It should be noted that not all researchers have found that recessions are more volatile than recoveries in all countries. Mejia-Reyes (2000) finds that output has greater volatility in recessions for four out of eight Latin American countries surveyed (Bolivia, Chile, Mexico and Peru), but in the case of Colombia, expansions are more volatile than recessions. In three cases (Argentina, Brazil, and Venezuela) there appeared to be no difference in volatility across business cycle phases.

Upon obtaining the FTP results, we then estimated the probabilities in the probit model, as mentioned:

$$F^{-1}\big(P(S_t = 1)\big) = \tau_0 + \tau_1 D_{99} + \tau_2 y_{GF} + \tau_3 D_{99} \cdot y_{GF} + \tau_4 y_{US} + \tau_5 D_{99} \cdot y_{US} + \epsilon_t$$

Some may suggest that we should include these five variables in the original FTP equation. Should we do that, for some countries, MLE convergence is difficult to achieve. As mentioned before, in some cases, even when the convergence is achieved, the Hessian matrix fails to be positive definite at the point of supposed convergence; and standard errors of some estimates could not be obtained. In the current FTP/probit model, we assume that these five variables are second stage variables, which are included in the error term. The error term drives the economy to be in a recession/ an expansion. As the FTP MLE estimates are obtained, we are able to calculate the probability of S_t=1 or S_t=0 (i.e., being in a recession or an expansion) for each time period[14]. Thus, instead of incorporating all these variables in the original

[13] Unlike α_1, h_1 does not measure the differences between regime 0 and 1.

[14] Given $\sum_{S_t=0}^{1}\sum_{S_{t-1}=0}^{1}\sum_{S_{t-2}=0}^{1} f(\Delta y_t \mid S_t, S_{t-1}, S_{t-2}, \Phi_{t-1}) = f(\Delta y_t \mid \Phi_{t-1})$, we can update the probability by

FTP equation, we examine the direct impacts of these five variables on the peripheral nations' probabilities of being in recession (S_t=1), with Greece being the exception (state 0 is the recession state for Greece). As an illustration of the MS results, Figures 1 and 2 show the actual industrial production (IP) growth rates (the solid line) and the probability of being in recession (S_t=1) (the line with "+") for Ireland and Switzerland. The left vertical axis denotes the IP growth rate while the right vertical axis indicates the probability.

The results of the probit model are shown in Table 4. As displayed, for Greece and Denmark the null of no US impact could not be rejected, and thus the two US variables were not included in the probit estimation. Ireland and Spain, in contrast, both appear to have become less synchronized with the United States since the Euro's creation (estimates of τ_5 are negative and significant for both countries). Of course, our key parameter of interest is τ_3. The estimate of τ_3 is insignificant for Portugal and Spain, indicating that joining the Euro has led to no change in BCS with the center for these two nations. However, the estimate is positive and significant for Ireland, and negative and significant for Greece. Both of these estimates indicate a *decrease* in BCS for Ireland and Greece. Again, the estimate of α_1 is negative for Ireland, so an increase in the probability of being in regime 1 means an increase in the probability of recession; for Greece, the estimate of α_1 is positive, and thus a decrease in the probability of being in this regime is a decrease in the probability of being in an expansion.

In the cases of the three non-Euro countries, the estimate of τ_3 is insignificant for two-Sweden and Switzerland. Thus, failing to join the Euro has not led to a loss of BCS with the center of the Euro zone. However, for the case of Denmark, the coefficient is negative and significant, so an increase in central Euro-zone output growth since the creation of the euro now has a significantly larger effect in lowering the probability of a Danish recession.

$f(\Delta y_t \mid S_t, S_{t-1}, S_{t-2}, \Phi_{t-1}) / f(\Delta y_t \mid \Phi_{t-1}) = f(S_t, S_{t-1}, S_{t-2} \mid \Delta y_t, \Phi_{t-1}) = f(S_t, S_{t-1}, S_{t-2} \mid \Phi_t)$. Then integrating out S_{t-1} and S_{t-2}, we can get $P(S_t = 1|\Phi_t)$.

Once again, joining the Euro has led to no increase in BCS with respect to the center for any of the four Euro zone nations, and in two cases led to a decrease in BCS. Failing to join the Euro has led to no loss of BCS, and in one case, BCS with the center of the Euro zone increased despite (or maybe even because) of retaining one's historical currency.

The results from the TVTP and FTP/Probit exercises are not identical, although they are similar. While different Euro countries have seen a decrease in BCS depending on which method is used (Ireland and Portugal using the TVTP, Ireland and Greece with the FTP/Probit), and two different non-Euro nations have exhibited an increase in BCS (Sweden with the TVTP, Denmark with the FTP/Probit), the results are still fairly close in an important respect. For both methods, there was no case of a Euro country experiencing an increase in BCS with France and Germany, and there were two cases in which there was a significant decrease (and Ireland's decrease in BCS was significant with both estimators). Moreover, for both methods, there was no loss of BCS for any of the three non-Euro countries. Finally, with both techniques, one non-Euro zone member experienced an increase in BCS while retaining its own currency.

As a concluding exercise for the FTP model, another way to evaluate the impact of changes on output dynamics in these economies is to inspect the changes in probability directly. Table 5 shows the effects of the central euro-zone and U.S. economies on the probability of being in regime 1 (i.e., the recession regime for all countries except Greece). Using the mean values of the independent variables as benchmarks for both before and after 1999Q1, we calculate the change in the probability P(S_t=1) due to a one standard deviation difference in the German/French IP growth and/or the U.S. IP growth. The results can be classified into three groups. First, for Portugal, Spain, Sweden and Switzerland, column four of Table 5[15] shows that a one standard deviation increase in both German/French IP growth and U.S. IP

[15] In Table 5, we reported $\hat{F}(\hat{z}_2) - \hat{F}(\hat{z}_1) = \hat{F}(x_2'\hat{\beta}) - \hat{F}(x_1'\hat{\beta})$ where $\hat{z}_2$ and $\hat{z}_1$ are the predicted values derived from the probit model with $\hat{z}_2$ being the result from the explanatory variables (e.g., y_{GF} etc.) that are 0.5 standard

growth reduces the probability of being in the recession (and more volatile) regime by 14.2%, 2.6%, 5.2%, 20.2% (respectively) before 1999Q1 and 22.7%, 11.8%, 2.8%, 43.6% after 1999Q1. Columns five and six show that, for Portugal and Switzerland, the reduction of 22.7% and 43.6% after 1999Q1 is primarily due to the increase of U.S. IP growth rate; not so much by the increase of German-French output growth. Secondly, for Greece and Denmark, a one standard deviation increase in both German-French and U.S. IP growth increase the probability of being in regime 1 by 4.9% and 0.8% before 1999Q1 and decrease the probability of being in regime 1 by 9.5% and 22.1% after 1999Q1. To repeat, note that regime 1 is the contraction regime for Denmark while it is an expansion regime for Greece. Third, for Ireland, a one standard deviation increase in both German-French and U.S. IP growth increases the probability of being in the recession regime both before and after 1999Q1 by 8% and 9%. But the decomposition in columns five and six shows that, after 1999Q1, U.S. economic expansion actually decreases Ireland's probability of being in recession by 18.2% but German/French expansion increases that probability by 27.3%.

In terms of the Euro's impact on BCS, we examine the impact of French-German growth on the probability of being in regime 1 since the currency's introduction in 1999. For the four Euro-zone countries, there is no significant impact in two cases (Portugal and Spain). There is a positive effect, significant at the five percent level for Ireland, and a negative effect, significant at the ten percent level, for Greece. Since regime 1 is the expansion state for Greece, both Ireland and Greece appear to have become less synchronized with the central Euro economies since joining the EMU.

For the three non-Euro nations, there is no significant change in the impact of German-French output since the Euro's creation in two cases (Sweden and Switzerland). There is a negative impact,

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deviations higher than their respective mean values and $\hat{z}_1$ being the result from the explanatory variables (e.g., y_{GF} etc.) that are 0.5 standard deviations lower than the mean values. To calculate the standard error (or the *p*-value) of the reported values, note that $\hat{F}(\hat{z}_2) - \hat{F}(\hat{z}_1) = \hat{F}(x_2'\hat{\beta}) - \hat{F}(x_1'\hat{\beta}) = g(\hat{\beta})$. The delta method (Greene, 2008:1055-1056) yields approximate errors.

significant at the one percent level, on Denmark's probability of being in recession, resulting from center output growth in the post-Euro era. Thus, again, joining the Euro has led to no increase in BCS with France and Germany for four Euro nations, and a decrease in two of these countries. In contrast, failure to join the common currency has led to no decrease in BCS for three non-Euro nations, and has been followed by an increase in BCS in one case. All of these results are again the precise opposite of what the endogenous OCA theory would predict.

Conclusion

The debate over readiness for entry into the Euro zone has been split, going back prior to the Euro's creation, between those believing that prospective members should fulfill a number of criterion prior to joining, and those who thought the very act of joining would create the conditions necessary for the common currency's benefits to exceed costs. Eichengreen (2002) provides an excellent synopsis of the debate between the two camps. He concludes that theory and evidence point to the importance of reforming prior to the adoption of a common currency, and not expecting the act of joining to produce the desired economic changes.

Frankel and Rose, on the other hand, have presented much evidence, mostly based on data prior to the Euro's existence, that the act of joining the Euro, whatever its other effects, should increase business cycle synchronization among member nations. The data we have examined here on the smaller Euro (and non-Euro) countries yields no such hope. There has been no increase in synchronization between the large continental economies and the four smaller Euro members-indeed there has been divergence in two cases. Moreover, staying out of the Euro zone has not led to a decrease in synchronization between the central Euro economies and three smaller non-Euro economies-indeed in one, such synchronization has actually increased, despite (or perhaps because of) retaining the historic currency. These results are clearly the opposite of what would be expected by the endogenous OCA literature.

Despite the turmoil suffered in the Euro zone in the wake of the 2008-09 global economic downturn, especially among the smaller members of the EMU, accession to the Euro is still a strong aspiration among many Eastern European economies. There may indeed be benefits-possibly greater trade and access to better-developed capital markets (although estimates of the magnitude of such benefits vary widely). Whatever benefits the common currency may bring, evidence presented here indicates that entry to the EMU will not increase cyclical convergence, and thus one monetary policy for the mature Euro zone members and the transition economies could prove very problematic.

References

Alesina, A. and R. Barro (2002) "Currency Unions", *Quarterly Journal of Economics*, 117, 409-436.

Alesina, A., R. Barro and S. Tenreyro (2002) "Optimal Currency Areas", in *NBER Macroeconomics Annual 2002, 17,* pp. 301-356.

Barro R. and S. Tenreyro (2007) "Economic Effects of Currency Unions", *Economic Inquiry,* 45, 1-23.

Buscher, H. and H. Gabrisch (2011) "The Synchronization of Wage Dynamics Across the EMU Members: A Test of the Endogeneity Hypothesis", *Empirica*, forthcoming.

Economist (2005) "Traffic Lights on the Blink", August 18th, p. 18.

Eichengreen, B. (2001) "Capital Account Liberalization: What do the Cross-Country Studies Tell Us?", *World Bank Economic Review,*

Eichengreen, B. (2002) "When to Dollarize", *Journal of Money, Credit and Banking*, 34, 1-24.

Eichengreen, B. and R. Hausmann (1999) "Exchange Rates and Financial Fragility", *Proceedings, Federal Reserve Bank of Kansas City,* pp. 329-368.

Enders, W. (2010) Applied Econometric Time Series, Wiley.

Filardo, A. (1994) "Business Cycle Phases and Their Transitional Dynamics", *Journal of Business and Economic Statistics*, 12, 299-308.

Frankel and Rose (1996) "Economic Structure and the Decision to Adopt a Common Currency", *Seminar Papers 611,* Stockholm University, Institute for International Economic Studies.

Frankel, J. and A. Rose (1998) "The Endogeneity of the Optimum Currency Criteria", *Economic Journal*, 108, pp. 1009-1025.

Kim, Chang-Jin and Charles Nelson (1999) State Space Models with Regime Switching, MIT Press, Cambridge, USA.

Mejia-Reyes, P. (2000) "Asymmetries and Common Cycles in Latin America: Evidence from Markov-Switching Models", *Economia Mexicana, Nueva Epoca,* 9, 189-225.

Mitchell, W. (1927) *Busienss Cycles: The Problem and its Setting*, New York, National Bureau of Economic Research.

Mundell, R. (1961) “ A Theory of Optimum Currency Areas”, *American Economic Review,* 51, 657-665.

Persson, T. (2001) “Currency Unions and Trade: How Large is the Treatment Effect?”, *Economic Policy,* 33, 435-448.

Rose, A. (2008) “Is EMU Becoming an Optimum Currency Area? The Evidence on Trade and Business Cycle Synchronization”, *Working Paper, available at http://faculty.haas.berkeley.edu/arose/RecRes.htm.*

Santos, A. (2002) “Are Mexican Business Cycles Asymmetrical?”, *IMF Working Paper No. 02/150.*

Sosa, S. (2008) “External Shocks and Business Cycle Fluctuations in Mexico: How Important Are U.S. Factors?”, *IMF Working Paper WP/08/100.*

Von Hagen, J. and M. Neumann (1994) “Real Exchange Rates Within and Between Currency Areas: How Far Away is EMU?”, *The Review of Economics and Statistics,* 76, 236-244.

Willett, T., O. Permpoon and C. Wihlborg (2010) “Endogenous OCA Analysis in the Early Euro Experience”, *World Economy,* 33, 851-872.

Wolf, N. and A. Ritschl (2011) “Endogeneity of Currency Areas and Trade Blocs: Evidence from a Natural Experiment”, *Kyklos*, 64, 291-312.

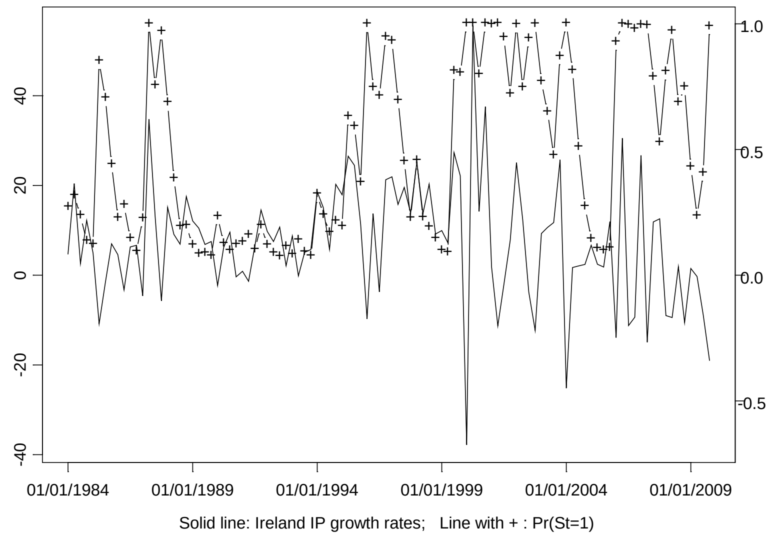

Figure 1: Ireland IP growth rates and $P(S_t=1)$

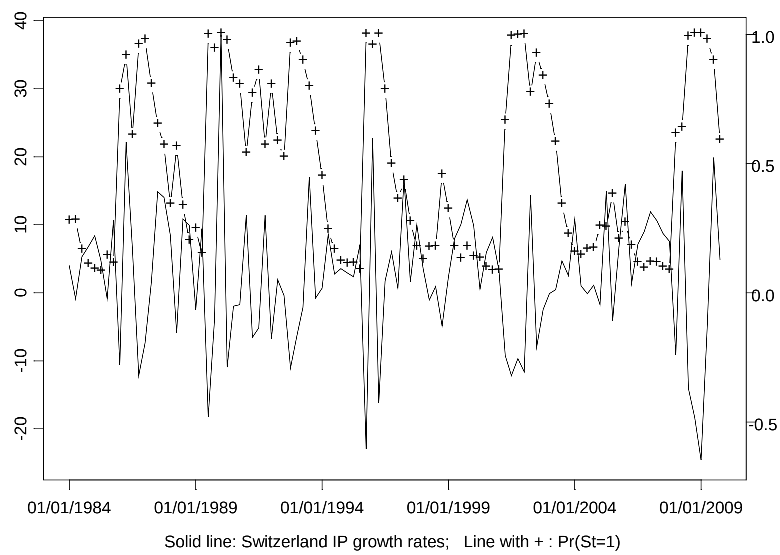

Figure 2: Switzerland IP growth rates and $P(S_t=1)$

Table 1: Estimates of Simple Regressions

	EU-Zone Members				Non-EU-Zone Members		
Estimates	Portugal	Ireland	Greece	Spain	Denmark	Sweden	Swiz
δ_1	-0.002 (0.277)	0.010 (0.039)	-0.002 (0.503)	-0.002 (0.230)	-0.000 (0.928)	-0.000 (0.910)	0.005 (0.077)
δ_2	0.453 (0.011)	0.070 (0.841)	0.272 (0.196)	0.509 (0.000)	1.534 (0.001)	0.350 (0.155)	1.000 (0.000)
δ_3	0.305 (0.257)	0.335 (0.529)	0.189 (0.550)	0.134 (0.536)	-1.099 (0.124)	0.678 (0.069)	-0.491 (0.122)
δ_4	0.271 (0.109)	1.464 (0.000)	-0.063 (0.749)	0.212 (0.122)	0.202 (0.650)	0.432 (0.065)	0.085 (0.666)
δ_5	-0.726 (0.038)	-1.465 (0.035)	0.079 (0.846)	0.160 (0.567)	0.474 (0.604)	-0.383 (0.423)	0.443 (0.278)
p-value of *H_0*: $\delta_2 = \delta_3 = 0$	(0.000)	(0.586)	(0.069)	(0.000)	(0.004)	(0.000)	(0.000)
p-value of *H_0*: $\delta_4 = \delta_5 = 0$	(0.092)	(0.000)	(0.949)	(0.097)	(0.630)	(0.179)	(0.305)
p-value of *H_0*: $\delta_2 = \delta_3 = \delta_4 = \delta_5 = 0$	(0.000)	(0.000)	(0.015)	(0.000)	(0.001)	(0.000)	(0.000)

Table 2: Parameter Estimates of TVTP MS Model, for both GF and D99*GF

	EU-Zone Members				Non-EU-Zone Members		
Estimates	Portugal	Ireland	Greece	Spain	Denmark	Sweden	Swiz
α_0	8.665 (0.000)	6.321 (0.002)	-0.346 (0.659)	1.497 (0.002)	2.237 (0.130)	1.992 (0.012)	9.586 (0.000)
α_1	-9.475 (0.000)	-0.251 (0.919)	5.344 (0.380)	-20.527 (0.000)	-3.902 (0.521)	-12.383 (0.223)	-14.373 (0.000)
ϕ_0	-0.824 (0.000)	-0.886 (0.000)	-1.044 (0.000)	-0.692 (0.000)	-1.227 (0.000)	-1.029 (0.000)	-1.270 (0.000)
γ_1	-0.129 (0.115)	-0.218 (0.159)	-0.051 (0.670)	-0.104 (0.247)	-0.177 (0.129)	-0.044 (0.597)	0.169 (0.172)
γ_2	-0.074 (0.378)	-0.089 (0.366)	-0.060 (0.463)	0.207 (0.008)	-0.234 (0.001)	-0.034 (0.569)	-0.066 (0.528)
h_0	4.345 (0.061)	42.159 (0.043)	40.315 (0.000)	19.953 (0.000)	90.455 (0.000)	48.516 (0.000)	44.660 (0.007)
h_1	53.215 (0.000)	257.747 (0.000)	331.776 (0.078)	38.143 (0.225)	795.352 (0.003)	882.308 (0.028)	77.081 (0.002)
$\theta_{\mathrm{p}0}$	2.849 (0.853)	5.708 (0.002)	-12.909 (0.043)	0.966 (0.158)	1.410 (0.180)	3.260 (0.146)	-1.580 (0.483)
$\theta_{\mathrm{p}1}$	-1.618 (0.106)	-2.500 (0.004)	20.0 (0.015)	0.205 (0.419)	-1.804 (0.244)	2.556 (0.015)	-2.932 (0.200)
$\theta_{\mathrm{p}2}$	2.241 (0.033)	3.009 (0.000)	-11.511 (0.699)	0.005 (0.956)	-20.0 (0.288)	-1.841 (0.002)	-2.846 (0.428)
$\theta_{\mathrm{q}0}$	1.221 (0.259)	6.409 (0.050)	2.580 (0.031)	4.910 (0.000)	2.011 (0.001)	4.769 (0.053)	-1.479 (0.337)
$\theta_{\mathrm{q}1}$	-0.342 (0.726)	-6.633 (0.168)	-1.488 (0.359)	0.048 (0.979)	0.141 (0.834)	3.594 (0.173)	1.289 (0.503)
$\theta_{\mathrm{q}2}$	-6.337 (0.580)	-12.587 (0.501)	3.041 (0.172)	-1.932 (0.446)	1.016 (0.506)	0.435 (0.824)	10.136 (0.101)
L-Like	-335.48	-400.04	-348.54	-300.02	-411.40	-359.00	-361.58
p-value of H_0: $\theta_{\mathrm{p}1} = \theta_{\mathrm{p}2} = \theta_{\mathrm{q}1} = \theta_{\mathrm{q}2} = 0$	(0.138)	(0.124)	(0.039)	(0.821)	(0.192)	(0.819)	(0.007)

- Numbers in the parentheses are *p*-values.

Table 3: Estimates of FTP Markov Switching Parameters

	EU-Zone Members				Non-EU-Zone Members		
Estimates	Portugal	Ireland	Greece	Spain	Denmark	Sweden	Swiz
α_0	1.107 (0.232)	10.286 (0.000)	0.306 (0.717)	0.992 (0.023)	3.774 (0.000)	2.586 (0.003)	5.668 (0.001)
α_1	-1.507 (0.568)	-3.912 (0.051)	0.232 (0.963)	-0.705 (0.395)	-2.420 (0.187)	-52.365 (0.008)	-6.236 (0.050)
ϕ_0	-0.439 (0.007)	-0.946 (0.000)	-1.135 (0.000)	-0.473 (0.000)	-1.183 (0.000)	-0.972 (0.000)	-1.123 (0.000)
γ_1	-0.472 (0.003)	-0.207 (0.132)	-0.017 (0.891)	-0.082 (0.447)	-0.171 (0.000)	-0.090 (0.296)	0.032 (0.865)
γ_2	-0.262 (0.104)	-0.108 (0.156)	-0.058 (0.483)	0.227 (0.009)	-0.271 (0.000)	-0.040 (0.506)	-0.018 (0.872)
h_0	23.458 (0.097)	9.058 (0.068)	35.631 (0.001)	2.669 (0.005)	0.010 (0.112)	57.194 (0.000)	31.238 (0.001)
h_1	124.579 (0.071)	217.455 (0.000)	267.194 (0.076)	41.587 (0.000)	323.356 (0.000)	272.011 (0.565)	158.586 (0.000)
$\tilde{p}$	0.754 (0.290)	0.962 (0.382)	0.456 (0.009)	0.982 (0.400)	0.978 (0.209)	0.288 (0.011)	0.869 (0.106)
$\tilde{q}$	0.868 (0.197)	0.690 (0.188)	0.891 (0.295)	0.919 (0.286)	0.634 (0.091)	0.975 (0.318)	0.885 (0.053)
Nob	103	103	103	103	103	103	103
L-Like	-353.21	-419.40	-362.98	-320.82	-429.00	-373.26	-380.09

Table 4: Estimates of Probit Model

	EU-Zone Members				Non-EU-Zone Members		
Estimates	Portugal	Ireland	Greece	Spain	Denmark	Sweden	Swiz
τ_1	0.038 (0.857)	1.790 (0.000)	-0.208 (0.233)	0.218 (0.379)	-0.368 (0.091)	-0.947 (0.008)	-0.537 (0.068)
τ_2	-0.032 (0.152)	-0.017 (0.304)	0.026 (0.112)	-0.039 (0.220)	0.003 (0.859)	0.044 (0.165)	-0.022 (0.513)
τ_3	0.037 (0.306)	0.150 (0.002)	-0.072 (0.008)	0.015 (0.819)	-0.091 (0.001)	-0.014 (0.793)	-0.004 (0.952)
τ_4	-0.046 (0.111)	0.088 (0.001)		-0.008 (0.686)		-0.128 (0.011)	-0.093 (0.007)
τ_5	-0.061 (0.137)	-0.240 (0.006)		-0.128 (0.030)		0.058 (0.536)	-0.065 (0.306)
τ_0	-0.253 (0.083)	-0.733 (0.000)	-0.852 (0.000)	-1.783 (0.000)	-0.489 (0.007)	0.038 (0.896)	0.594 (0.014)
p-value of H_0: $\tau_4 = \tau_5 = 0$	(0.001)	(0.001)	(0.308)	(0.047)	(0.593)	(0.026)	(0.001)

- Numbers in the parentheses are *p*-values.

Table 5: Impacts of Central Euro IP or U.S. IP on $P(S_t=1)$

			One SD Change in both y_{GF} and y_{US}	One SD Change in y_{GF} alone	One SD Change in y_{US} alone
EU-Zone	Portugal	$\Delta P(S_t=1)$ before 1999Q1	-0.142 (0.008)	-0.077 (0.157)	-0.065 (0.073)
		$\Delta P(S_t=1)$ after 1999Q1	-0.227 (0.005)	-0.031 (0.680)	-0.197 (0.002)
	Ireland	$\Delta P(S_t=1)$ before 1999Q1	0.08 (0.147)	-0.03 (0.544)	0.110 (0.005)
		$\Delta P(S_t=1)$ after 1999Q1	0.09 (0.078)	0.273 (0.010)	-0.182 (0.108)
	Greece	$\Delta P(S_t=1)$ before 1999Q1	0.049 (0.159)	0.049 (0.159)	
		$\Delta P(S_t=1)$ after 1999Q1	-0.095 (0.059)	-0.095 (0.059)	
	Spain	$\Delta P(S_t=1)$ before 1999Q1	-0.026 (0.084)	-0.026 (0.095)	-0.000 (0.993)
		$\Delta P(S_t=1)$ after 1999Q1	-0.118 (0.117)	-0.036 (0.436)	-0.078 (0.149)
Non EU-Zone	Denmark	$\Delta P(S_t=1)$ before 1999Q1	0.008 (0.858)	0.008 (0.858)	
		$\Delta P(S_t=1)$ after 1999Q1	-0.221 (0.000)	-0.221 (0.000)	
	Sweden	$\Delta P(S_t=1)$ before 1999Q1	-0.052 (0.556)	0.138 (0.086)	-0.189 (0.005)
		$\Delta P(S_t=1)$ after 1999Q1	-0.028 (0.497)	0.013 (0.869)	-0.041 (0.691)
	Switz	$\Delta P(S_t=1)$ before 1999Q1	-0.202 (0.000)	-0.098 (0.131)	-0.105 (0.023)
		$\Delta P(S_t=1)$ after 1999Q1	-0.436 (0.000)	-0.107 (0.497)	-0.341 (0.001)