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Did the Euro Common Currency Increase or Decrease Business Cycle Synchronization for its Member Countries?

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We use two variants of Markov switching models to assess changes in output synchronization since the creation of the euro. Out of eight eurozone countries investigated, only one—the Netherlands—has synchronization increased since euro adoption, supporting the 'endogenous optimal currency area' argument of Frankel and Rose. However, in three other cases, business cycle synchronization actually *fell* since the euro's creation. Thus the 'endogeneity' of the optimal currency area criteria can go both ways—adopting a common currency may increase synchronization for nations ready for a common currency, but it can lower synchronization for nations that are far from synchronized before monetary unification.

INTRODUCTION

Preceding its beginning in 1999, the eurozone attracted sceptics 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 benefited from a tighter monetary policy over 1999–2006. In contrast, wealthier 'centre' nations such as France and Germany experienced very sluggish growth in the initial years of the currency and may well have benefited from looser money and lower interest rates than the European Central Bank (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 seem to be the standard prescription. Yet at the same time, the ECB was criticized, at least until its actions in 2012, for being too tight in the face of the recent 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—resigned in 2011 in apparent protest against central bank attempts to alleviate the debt crisis by buying sovereign bonds.) The eurozone, in short, may not truly be 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 centre and peripheral countries of the eurozone may be one of the most important reasons for doubts that the now nineteen-member union is truly an OCA.

On the other hand, a growing academic literature over the last decade (see Rose (2008)[1] for a summary and meta-analysis) has produced results indicating that even if a currency union (CU) such as the eurozone is not an OCA 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. Second, the increase in trade will increase business cycle synchronization (BCS) 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, on 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.

In addition to trade and BCS, labour mobility may also be endogenous to the existence of a CU. Eichengreen (2002) points out that in the absence of nominal exchange rate depreciation, greater labour market flexibility is important to the functioning of a CU. He further notes that some claim that the very act of joining a CU may 'produce the necessary labour market reforms since unions will recognize the absence of alternative adjustment mechanisms' (Eichengreen 2002, p. 2).

However, endogeneity could go in both directions—that is, there are channels through which a common currency may lower rather than increase BCS. To take one example, a currency union may increase foreign borrowing (see Eichengreen and Hausmann (1999) on exchange rate regimes and capital flows) as capital may flow from larger countries into smaller, peripheral nations, fuelling growth (and perhaps bubbles) in the latter, and may lower rather than raise synchronization. Another mechanism through which a common currency may lower BCS between members is different inflation rates in different 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' (*The Economist*, 2005). There are thus a number of ways in which a currency union could increase or decrease output comovements.

The impact of monetary integration on BCS is therefore an empirical issue. And indeed there have been many previous empirical papers on the topic (see Rose (2008) for an excellent survey). In these papers, BCS is quite often gauged with the correlation of de-trended income between two countries. It is highly questionable whether this technique gives an accurate measure of BCS. Mink *et al*. (2012) note that output gaps between countries may display a high correlation, but that business cycles could still be very different due to wide variation in output gap magnitudes. This could very well be the case when gauging the impact of large, 'centre' economies on smaller, 'peripheral' countries.

Accordingly, we take a different approach. We employ a technique that directly estimates the impact of 'centre' output on 'peripheral' business cycles. This Markov switching (MS) technique has been employed elsewhere to model business cycles and examine the impact of cyclical shocks between countries. Both the time-varying transitional probability (TVTP) and the traditional fixed transitional probability (FTP) MS models are used to examine the output dynamics of these small economies. The TVTP model assumes that certain economic variables, such as the centre nations' output growth, affect the transitional probability, that is, the probability of a transition from expansion (contraction) to contraction (expansion). In contrast, the FTP model assumes that the transitional probabilities are exogenous and hence the probabilities are not modelled as functions of economic variables. In an FTP model, the unobserved state variable, dictated by constant probabilities, determines the state of the economy, that is, being in 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 centre nations' output growth directly into the original FTP equation, we use a probit model to investigate the impacts of the centre nations' output growth on the peripheral nation's probability of being in recession.[2] What we find is that in one case—the Netherlands—in one of these two variants of MS models employed, there has been an increase in synchronization with the centre nations of France and Germany since euro adoption. However, in two other cases—Ireland and Greece in FTP, and Ireland and Portugal in TVTP—there has been a decrease in synchronization since the euro's creation.

Our results thus indicate that a monetary union may be followed by more or by less synchronization among member nations. We also take note of the following in our interpretation. Several previous studies (von Hagen and Neumann (1994); Mink *et al.* (2012)) indicate that the Netherlands seemed fairly 'ready' for the euro before monetary union took place. In contrast, metrics applied to Greece, Ireland and Portugal in Mink *et al.* (2012) suggest a lower level of readiness for a common currency. It is thus possible that the act of joining a common currency may have some impact on BCS—it could be positive if a country appears fairly integrated with other members before joining, but it could well be also negative for nations that join with low levels of BCS prior to the actual currency adoption.

This paper proceeds as follows: the next section details the literature on OCAs and their possible endogeneity; Section II discusses the methodology, and Section III presents our results; Section IV concludes.

I. Previous Literature

Mundell's 1961 paper is cited as the beginning of the OCA literature. Important early contributions were also made by McKinnon (1963) and Kenen (1969). Criteria for whether a group of countries should join a common currency included the extent of trade and the mobility of labour (the greater each was, the more likely countries would form an OCA). A system of fiscal transfers was later added. More recent attention has turned to the synchronization of business cycles as an OCA criterion—with one currency, there is one central bank and only one monetary policy, which could be very difficult if different countries in a monetary union are in different phases of the business cycle at a given time.

Given the different sizes and structures of the varied eurozone economies, business cycle synchronization was a major concern prior to the euro's existence. It has been posited, however, that such concerns may be overblown, in that trade and BCS may be endogenous to the very act of joining a CU. The findings of Frankel and Rose (1996) indicate that creating a CU will bolster the extent of trade among members (the authors' initial results suggested a tripling of trade as a consequence of monetary union) and moreover that trade would increase BCS. At first glance, the impact of greater trade, if the models of Ricardo or Hecksher and Ohlin are a guide, could cause more specialization, and therefore a less similar response to shocks (Krugman (1993)). Frankel and Rose (1998, p. 1010), however, claim that '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'.

The conjectures of Frankel and Rose, if correct, bolster the case for a CU, even if the potential members do not appear to have high levels of BCS prior to joining the monetary union. The CU will greatly expand trade, and then the greater trade will increase BCS. Thus nations that do not appear ready for one money can become ready by the act of creating the common currency. The authors state in their 1996 study on whether Sweden should join the coming Eurozone: '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).

It is not at all clear, however, that the hypothetical impact of joining a common currency should obviate all concern about BCS before joining. First, even if there is some synchronizing impact from monetary integration, the magnitude of this effect may not be sufficient to make a group of countries a truly *optimal* currency area. Rose (2008, p. 7) himself states that although his meta-analysis seems to show that a common currency does increase BCS, 'whether this effect is big enough to make Europe an OCA remains to be seen'.

Moreover, there are concerns over whether the cross-country studies on which claims of a synchronizing effect from CUs are based might be plagued with endogeneity bias. Eichengreen (2001) explains in detail how endogeneity is a serious problem for cross-country studies of macroeconomic policies. In particular, a country that adopts a certain macroeconomic policy may have other policy 'imbalances' or attributes, some of which will be difficult to observe, and the estimates of the effect of the macroeconomic policy will be biased by hard-to-observe imbalances and other factors. For instance, in the context of capital controls, Eichengreen points to the many estimates of the impact of such controls, and notes that the '"effects" will reflect the influence of these deeper policy problems as much as those of the capital controls themselves' (Eichengreen 2001, p. 344). In the context of the present study, trying to gauge the impact of a CU on synchronization with a cross-country dataset may lead to biased estimates due to hard-to-observe factors such as political orientation, trade patterns, institutions and anything else that may affect a country's propensity to join a CU and the level of synchronization with other CU members. These endogeneity problems make Markov switching analysis, which we perform in this study, more appealing.

This simultaneity bias does appear to be an issue for studies on the 'endogeneity' of the OCA criteria. Persson (2001) expresses doubts about the very large estimates of the impact of CUs on trade, and, using a matching estimator, finds no statistically significant effect of CUs on trade. Wolf and Ritschl (2011) estimate the impact of 1930s common currencies on trade, and find that standard results suggest very large trade effects. However, the authors investigate further and find that the results were affected by pervasive endogeneity. They caution against buying into high 'trade creation effects' from standard empirical models of CUs and trade. Havranek (2010) employs a meta-analysis of the effect of the euro itself on trade, and finds, after correcting for certain bias, that there has been no significant effect of this particular CU on trade.

Another major issue is the second link in the endogenous OCA hypothesis—even if a CU increases trade, how much (if at all) does trade boost BCS? Of those 20 papers about the trade–BCS linkage in the meta-analysis of Rose (2008), 14 papers have significantly positive results, while six show insignificant linkage between trade and BCS. However, in addition to this specific linkage (currency unions raise trade, and trade raises business cycle synchronization), it may be that 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 Hausmann (1999, pp. 91–2) for a discussion on exchange rate rigidity and foreign borrowing), then capital may flow from relatively stagnant 'centre' countries to faster growing 'peripheral' nations with higher returns, fuelling growth (and perhaps bubbles) in the peripheral nations. This process, which would describe the experience of several smaller eurozone 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 *et al*. (2002) find that CUs do raise trade, but do not, in most specifications, have a significant impact on output comovement. Using a panel approach, and attempting to directly address the issue of endogeneity between trade and common currencies, Barro and Tenreyro (2007) find that trade is enhanced by CUs, while the comovement of output actually *decreases* in response to sharing the same currency.

Along these lines, Buscher and Gabrisch (2012) find that the euro has had little impact on synchronizing nominal wage dynamics in the eurozone, as asymmetries in nominal wage formation continue to persist across the continent's common currency area. In another scenario, Willett *et al*. (2010) examine the euro and BCS through the correlation of output and consumption, both within eurozone countries and within non-eurozone European nations. They find that the growth in the correlation of output was greater for European countries outside the eurozone than for those within the currency union!

In addition to generating higher capital flows, another way in which a CU could decrease rather than increase BCS is through different inflation rates in member countries. As there is only one monetary policy interest rate, low-inflation countries have higher real interest rates than do countries with higher inflation. Patrick Artus critically noted this phenomenon in *The Economist* in 2005, when German growth was sluggish but smaller countries in the periphery grew briskly.

Finally, there are some concerns about the techniques used to measure the BCS by some researchers. In the literature, BCS is quite often (but not always) measured as the correlation coefficient estimated from de-trended output levels between country i and country j. Some are not sure whether these correlation coefficients could really capture synchronization.[3] Mink *et al*. (2012) mention that a perfectly correlated output gap may still imply a large discrepancy in synchronization because of the large differences in cyclical magnitudes. Our paper investigates BCS in the Economic and Monetary Union (EMU) from a different angle. By using the data of representative smaller nations from both eurozone and non-eurozone countries, we compare the impacts of the larger economies of the eurozone on the transitional probabilities of the smaller nations. In this paper, an additional negative (positive) impact of the centre large economies on the small peripherals' probabilities of staying in recession after joining the euro would indicate an increase (decrease) in BCS.

II. Data and Methodology

Our goal is to determine how much, if at all, the euro has increased or decreased BCS between the larger and smaller economies of this most famous modern currency area. If the proposition of Frankel and Rose—that of an endogenous OCA in which BCS rises after a monetary union is formed—is valid, then we should observe an increase in BCS between the large and small EMU countries, and no such an increase (or at least a smaller increase) in BCS between centre EMU nations and small non-EMU countries. 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 use the USA 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 USA during the euro's existence.

The datasets 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 centre and the periphery as follows. The centre countries are Germany and France. We choose Germany and France as centre nations as they are the two largest economies in the monetary union. We then choose eight smaller euro economies. The first are the well-known PIIGS (Portugal, Italy, Ireland, Greece and Spain). All of these countries experienced serious difficulties after the 2008 global financial crisis. Italy was the largest of these, and avoided a bailout but still experienced a net current account deficit over the first decade of the euro, and faced rising interest rates on its government debt as investors got nervous in 2011. Ireland, Greece and Portugal, of course, suffered very severe financial crises and received bailouts from the 'troika'—the ECB, IMF and European Commission. Spain would agree to a bailout of its banking system in 2012. Of course, we do not wish to compare only centre and crisis countries. We accordingly include three smaller eurozone nations—Austria, Belgium and the Netherlands—that avoided sharp housing and banking crises. Finally, we want to include, for comparison, some smaller, non-euro countries to see how BCS evolved in the absence of a common currency. The peripheral non-EMU countries that we examine for comparison are Denmark, Switzerland and Sweden. Denmark and Sweden have both considered eurozone membership, and Switzerland has managed its monetary policy in response to the Swiss franc's movements against the euro. In addition, we include the USA. This is in an effort to control for external forces that may affect the estimates within Europe.

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

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

where y_t is the log change in peripheral country output; $D_{99} = 1$ if $t \geq$ 1999:Q1, while $D_{99} = 0$ if $t <$ 1999:Q1; 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, then centre output has increased its impact on the periphery since euro adoption. This would be evidence in favour 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 US output growth, and its interaction with the euro years, in this equation.

It is important to test for non-linearity. Though there are many procedures to test against certain particular non-linear 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 non-linear time series models. The idea behind this test is that if a given series is indeed linear, then the residuals from the linear model should be independent and identically distributed (iid). 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 dataset; we find that the null hypothesis of the data being iid is rejected for Austria, Belgium, Denmark, Greece, Ireland, Italy, the Netherlands, Portugal, Switzerland and Sweden. For Spain, if the epsilons are 1, 1.5 and 2 standard deviations of the dataset, then the null is rejected. The BDS test results thus indicate a call for a non-linear model.

Given the results of the non-linearity 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 non-linear process. Indeed, the literature on business cycles, going back to Mitchell (1927), holds that such cycles are clearly non-linear. Mitchell found that recessions exhibited more volatility than expansions; thus the dynamics of recoveries and slumps are different and should be modelled as such. To gather evidence in a more rigorous fashion, we employ Markov switching (MS) models, which, since Hamilton's (1989) pioneering paper, have been utilized to investigate business cycle fluctuations. (To anticipate our results, we will find, in our TVTP specification to be explained below, that the recession regime is more volatile than the expansion regime in most of the countries, which is in line with Mitchell's findings and highlights the importance of non-linear modelling.)

Before proceeding with describing our MS model, we note that one approach used in the past to understand business cycles in different regimes has been to attempt to adjust for the pattern of shocks—i.e. if business cycles over a given regime exhibit more or less volatility, cohesion, etc. due to a regime or due to a pattern of shocks. Some papers—such as Bayoumi and Eichengreen (1992)—attempt to identify the pattern of shocks with Blanchard–Quah decomposition in a vector autoregression (VAR); others use techniques such as Kalman filtering to identify different types and patterns in shocks. However, the identifying assumptions are highly questionable and thus may 'misidentify' the pattern of shocks. For instance, von Hagen and Neumann (1994) criticize the approach of Bayoumi and Eichengreen (1992), and state that the authors have misidentified shocks. Therefore, given these difficulties, we acknowledge that it would be optimal to find a way to identify the pattern of shocks, but given the likely failure of the techniques used to properly do so, we employ the MS model.

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

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

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

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

where $\rho = \sum_{i=1}^{p} \beta_i$ is a measure of persistence.[4]

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. Since the mean relates to the average output growth rate and the variance depicts the output growth volatility, our MS model allows different output volatility in two states of the economy, that is, expansion and contraction may have different volatilities. The previous model can thus be expressed as

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

where $\phi_0 = \rho - 1$. A two-state Markov process has transition probabilities

$$\Pr(S_t = 0|S_{t-1} = 0, \Phi_{t-1}) = \tilde{q}_t \quad \text{and} \quad \Pr(S_t = 1|S_{t-1} = 1, \Phi_{t-1}) = \tilde{p}_t$$

for the time-varying probability case, and

$$\Pr(S_t = 0|S_{t-1} = 0) = \tilde{q} \quad \text{and} \quad \Pr(S_t = 1|S_{t-1} = 1) = \tilde{p}$$

for the constant probability case. 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, then 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.[5] For the variance, $\mathrm{Var}(u_t) = h_0$ when $S_t = 0$, and $\mathrm{Var}(u_t) = h_1$ when $S_t = 1$. In this paper, we use $p = 3$.[6]

As mentioned above, 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 variables D_{99}, $y_{\mathrm{GF,t-1}}$, $D_{99}y_{\mathrm{GF,t-1}}$, $y_{\mathrm{US,t-1}}$, $D_{99}y_{\mathrm{US,t-1}}$ 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 each variable one at a time, the author found that only one of the candidate variables showed significance in the TVTP model.

Given the difficulties in obtaining meaningful results for some countries,[7] we follow two approaches. First, to obtain convergent estimates of a TVTP MS model, we drop the US variables ($y_{\mathrm{US,t-1}}$ and $D_{99}y_{\mathrm{US,t-1}}$) and the euro dummy ($D_{99}$) intercept as we cannot obtain convergence when they are included in the model. We then examine the impact, in the TVTP estimation, of $y_{\mathrm{GF,t-1}}$ and $D_{99}y_{\mathrm{GF,t-1}}$. Our second approach is 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 the TVTP model, we thus specify the transition probabilities as

$$\tilde{p}_t = \frac{\exp(\theta_{p0} + \theta_{p1} y_{\mathrm{GF},t-1} + \theta_{p2} D_{99} y_{\mathrm{GF},t-1})}{1 + \exp(\theta_{p0} + \theta_{p1} y_{\mathrm{GF},t-1} + \theta_{p2} D_{99} y_{\mathrm{GF},t-1})},$$
$$\tilde{q}_t = \frac{\exp(\theta_{q0} + \theta_{q1} y_{\mathrm{GF},t-1} + \theta_{q2} D_{99} y_{\mathrm{GF},t-1})}{1 + \exp(\theta_{q0} + \theta_{q1} y_{\mathrm{GF},t-1} + \theta_{q2} D_{99} y_{\mathrm{GF},t-1})}.$$

Thus for the TVTP model the parameters are (α_0, α_1, ϕ_0, γ_1, γ_2, h_0, h_1, θ_{p0}, θ_{p1}, θ_{p2}, θ_{q0}, θ_{q1}, θ_{q2}), and for the FTP model the parameters are $(\alpha_0, \alpha_1, \phi_0, \gamma_1, \gamma_2, h_0, h_1, \tilde{p}, \tilde{q})$. Maximum-likelihood estimation (MLE) is implemented to estimate the parameters. The log-likelihood function is

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

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

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

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 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 implement the following probit model for each of the eleven small nations:

$$F^{-1}(\Pr(S_t = 1)) = \tau_0 + \tau_1 D_{99} + \tau_2 y_{\text{GF}} + \tau_3 D_{99} \cdot y_{\text{GF}} + \tau_4 y_{\text{US}} + \tau_5 D_{99} \cdot y_{\text{US}} + \varepsilon_t.$$

As noted, y_{GF} and y_{US} are German–French and US industrial production growth rates, respectively, and D_{99} measures the post-1998 euro quarters. The coefficients τ_2 and τ_4 thus measure the association of German–French and US output growth, respectively, on the probability of recession[8] in the eleven smaller euro and non-euro countries. The interaction coefficients τ_3 and τ_5 are the additional association of German–French and US output growth on the probability of recession post euro adoption. If the Frankel and Rose theory of endogenous OCAs is correct, τ_3 should be negative and significant for the eight smaller euro economies. That is, since euro adoption, an increase in German–French output should lower the probability of recession in the smaller euro economies more than before the creation of the common currency.

III. Results

Table 1 contains results for the simple regressions. In six of the eight cases for eurozone members (Austria, the Netherlands, Portugal, Ireland, Greece and Spain), the estimate of δ_3— the interaction coefficient with the euro era and German–French output growth—is insignificant. These results indicate that based on simple linear regression, none of these six eurozone countries in the sample has experienced an increase in BCS with the centre of the monetary union since the euro was created. For two eurozone nations (Belgium and Italy), the estimated coefficient is positive and significant. At first glance, this might be interpreted to mean that in two of the eight (one-quarter) of the cases, there was a positive impact just from joining the EMU.

TABLE 1
ESTIMATES OF SIMPLE REGRESSIONS

	Eurozone members								Non-eurozone members		
Estimates	Austria	Belgium	Netherlands	Portugal	Ireland	Italy	Greece	Spain	Denmark	Sweden	Switzerland
δ_1	0.007	0.000	0.002	−0.002	0.010	−0.003	−0.002	−0.002	−0.000	−0.000	0.005
	(0.001)	(0.871)	(0.393)	(0.277)	(0.039)	(0.080)	(0.503)	(0.230)	(0.928)	(0.910)	(0.077)
δ_2	0.251	−0.360	0.229	0.453	0.070	0.188	0.272	0.509	1.534	0.350	1.000
	(0.030)	(0.03)	(0.122)	(0.011)	(0.841)	(0.086)	(0.196)	(0.000)	(0.001)	(0.155)	(0.000)
δ_3	0.252	1.088	−0.217	0.305	0.335	0.575	0.189	0.134	−1.099	0.678	−0.491
	(0.221)	(0.000)	(0.411)	(0.257)	(0.529)	(0.003)	(0.550)	(0.536)	(0.124)	(0.069)	(0.122)
δ_4	0.545	0.342	0.258	0.271	1.464	0.420	−0.063	0.212	0.202	0.432	0.085
	(0.000)	(0.047)	(0.145)	(0.109)	(0.000)	(0.001)	(0.749)	(0.122)	(0.650)	(0.065)	(0.666)
δ_5	−0.021	−0.134	0.641	−0.726	−1.465	−0.060	0.079	0.160	0.474	−0.383	0.443
	(0.942)	(0.724)	(0.104)	(0.038)	(0.035)	(0.834)	(0.846)	(0.567)	(0.604)	(0.423)	(0.278)
p-value of H_0:$\delta_2 = \delta_3 = 0$	(0.001)	(0.000)	(0.301)	(0.000)	(0.586)	(0.000)	(0.069)	(0.000)	(0.004)	(0.000)	(0.000)
p-value of H_0:$\delta_4 = \delta_5 = 0$	(0.000)	(0.115)	(0.015)	(0.092)	(0.000)	(0.002)	(0.949)	(0.097)	(0.630)	(0.179)	(0.305)
p-value of H_0:$\delta_2 = \delta_3 = \delta_4 = \delta_5 = 0$	(0.001)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.015)	(0.000)	(0.001)	(0.000)	(0.000)

Notes
Numbers in parentheses are p-values.

On the other hand, for the three *non*-eurozone countries, while the interaction term is insignificant in two of three nations (Denmark and Switzerland), it is positive and significant (at the 10% level) for Sweden. That is, despite failure to join the eurozone, Sweden's economy appears more affected by German–French output changes since the euro's creation. This is in contrast to six of the eight eurozone countries that display no such increase in BCS with the centre of the eurozone since adopting the common currency. Of course, given the findings of the BDS tests, none of these results from linear modelling should be accorded a high level of confidence.

It may be questionable to include US output growth in the model. We therefore test the restriction H_0:$\delta_4 = \delta_5 = 0$. For Austria, the Netherlands, Portugal, Ireland, Italy and Spain, our results indicate that the inclusion of US output growth is relevant, at least at the 10% level. Of perhaps greater interest, while six of the eight eurozone nations have not become more synchronized with the centre of the CU since the EMU began, four countries in the EMU—Austria, Belgium, Ireland and Italy—appear to have become more synchronized with the USA during this time!

We also test for H_0:$\delta_2 = \delta_3 = 0$. As displayed in Table 1, at the 10% significance level, all except Ireland and the Netherlands display a responsiveness to the eurozone centre economy. Finally, the test results of H_0:$\delta_2 = \delta_3 = \delta_4 = \delta_5 = 0$ are reported on the last row of Table 1, and as displayed, the null is rejected for all countries. But again, given the inherent non-linearity of output in these countries as revealed not just by theory and intuition but also the BDS tests, the results of these linear regressions are really nothing more than informal 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 except Belgium, regime 1 is more volatile than regime 0. As for the average IP growth rates, all countries, except Belgium and Greece, have negative estimates of α_1, that is, regime 1 is the contraction regime for all countries except Belgium and Greece. Thus for Belgium and Greece only, regime 1 is the expansion regime. Note that θ_{qi}, where $i = 1, 2, 3$, are related to regime 0, while θ_{pi} are related to regime 1. In particular, θ_{p2} measures the impact of the interaction of German–French output changes in the euro years on the probability of staying in recession for all countries except Belgium and 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 Belgium and Greece (where it measures the impact on the probability of staying in a recession). None of the estimates of θ_{q2} is significant. For eurozone members, the estimates of θ_{p2} are insignificant for Austria, Belgium, the Netherlands, Italy, Greece and Spain, which indicates no increase in BCS since joining the eurozone. Notably, however, θ_{p2} is positive and significant for both Portugal and Ireland. That is, an increase in German–French output, since the euro's introduction, raises the likelihood of a recession in these two eurozone nations—a result that is exactly the opposite of what would be expected based on the Frankel–Rose hypothesis on endogenous OCAs. Joining the euro has apparently led to less not more synchronization for these two smaller eurozone 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 eurozone's centre since the euro's introduction, despite Sweden's failure to join the eurozone. Thus none of the eight eurozone 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

TABLE 2
PARAMETER ESTIMATES OF TVTP MS MODEL, FOR BOTH GF AND D_{99}*GF

Estimates	Eurozone members								Non-eurozone members		
	Austria	Belgium	Netherlands	Portugal	Ireland	Italy	Greece	Spain	Denmark	Sweden	Switzerland
α_0	5.682	−3.959	4.014	8.665	6.321	1.119	−0.346	1.360	0.467	1.992	9.586
	(0.000)	(0.470)	(0.000)	(0.000)	(0.002)	(0.054)	(0.659)	(0.011)	(0.751)	(0.012)	(0.000)
α_1	−8.004	5.763	−7.546	−9.475	−0.251	−14.963	5.344	−10.412	1.753	−12.383	−14.373
	(0.001)	(0.302)	(0.005)	(0.000)	(0.919)	(0.127)	(0.380)	(0.061)	(0.854)	(0.223)	(0.000)
ϕ_0	−0.989	−1.107	−1.468	−0.824	−0.886	−0.767	−1.044	−0.599	−0.941	−1.029	−1.270
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
γ_1	−0.052	0.091	0.186	−0.129	−0.218	−0.168	−0.051	−0.123	−0.349	−0.044	0.169
	(0.677)	(0.373)	(0.171)	(0.115)	(0.159)	(0.154)	(0.670)	(0.227)	(0.011)	(0.597)	(0.172)
γ_2	0.124	0.078	0.002	−0.074	−0.089	−0.019	−0.060	0.227	−0.285	−0.034	−0.066
	(0.183)	(0.319)	(0.986)	(0.378)	(0.366)	(0.814)	(0.463)	(0.007)	(0.000)	(0.569)	(0.528)
h_0	27.312	351.008	32.582	4.345	42.159	27.271	40.315	18.577	123.370	48.516	44.660
	(0.001)	(0.021)	(0.000)	(0.061)	(0.043)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.007)
h_1	66.932	39.857	122.815	53.215	257.747	433.825	331.776	139.304	1209.583	882.308	77.081
	(0.001)	(0.000)	(0.002)	(0.000)	(0.000)	(0.118)	(0.078)	(0.079)	(0.022)	(0.028)	(0.002)
θ_{p0}	−0.184	6.951	0.379	2.849	5.708	−7.300	−12.909	−0.785	0.604	3.260	−1.580
	(0.912)	(0.037)	(0.676)	(0.853)	(0.002)	(0.585)	(0.043)	(0.922)	(0.546)	(0.146)	(0.483)
θ_{p1}	−1.706	4.841	8.386	−1.618	−2.500	−0.471	20.0	19.053	−1.360	2.556	−2.932
	(0.503)	(0.079)	(0.310)	(0.106)	(0.004)	(0.998)	(0.015)	(0.704)	(0.381)	(0.015)	(0.200)
θ_{p2}	−0.292	1.924	−9.606	2.241	3.009	−22.486	−11.511	−23.085	5.963	−1.841	−2.846
	(0.931)	(0.420)	(0.261)	(0.033)	(0.000)	(0.906)	(0.699)	(0.647)	(0.528)	(0.002)	(0.428)
θ_{q0}	0.664	6.037	15.513	1.221	6.409	6.866	2.580	3.954	2.475	4.769	−1.479
	(0.760)	(0.105)	(0.632)	(0.259)	(0.050)	(0.100)	(0.031)	(0.002)	(0.000)	(0.053)	(0.337)
θ_{q1}	21.421	2.285	40.997	−0.342	−6.633	0.160	−1.488	0.125	0.133	3.594	1.289
	(0.356)	(0.226)	(0.600)	(0.726)	(0.168)	(0.975)	(0.359)	(0.936)	(0.863)	(0.173)	(0.503)
θ_{q2}	−13.569	−1.004	7.810	−6.337	−12.587	7.145	3.041	−1.880	1.078	0.435	10.136
	(0.588)	(0.536)	(0.887)	(0.580)	(0.501)	(0.329)	(0.172)	(0.480)	(0.321)	(0.824)	(0.101)
Log-likelihood	−327.56	−343.47	−335.27	−335.48	−400.04	−316.46	−348.54	−301.38	−415.60	−359.00	−361.58
p-value of H_0:$\theta_{p1} = \theta_{p2} = \theta_{q1} = \theta_{q2} = 0$	(0.300)	(0.073)	(0.027)	(0.138)	(0.124)	(0.171)	(0.039)	(0.821)	(0.192)	(0.819)	(0.007)

Notes
Numbers in parentheses are p-values.

the three non-eurozone 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 quite opposite to those predicted by the Frankel and Rose 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 $\theta_{p1} = \theta_{p2} = \theta_{q1} = \theta_{q2} = 0$. 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,[9] the likelihood ratio test of $H_0{:}\theta_{p1} = \theta_{p2} = \theta_{q1} = \theta_{q2} = 0$ indicates that at the 5% significance level, the null hypothesis cannot be rejected for all countries, except the Netherlands,[10] Greece and Switzerland. Thus compared to the FTP, TVTP is actually not a clearly preferred model for Austria, Belgium, Portugal, Ireland, Italy, 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 economies' business cycles. So for a more robust investigation, we 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. 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.[11] Thus for the period 1983:Q1 to 2009:Q4 in Belgium, Italy and Sweden, there is a significant decline in the mean industrial production growth rate in regime 1 (i.e. when $S_t = 1$). The long-run average industrial growth rates in expansion for Austria, Belgium, the Netherlands, Portugal, Ireland, Italy, Greece, Spain, Denmark, Sweden and Switzerland are 4.69%, 1.45%, 2.01%, 2.52%, 8.29%, 1.48%, 0.570%, 2.10%, 3.19%, 2.66% and 5.05%, respectively, while the average growth rates in recession are 1.28%, −2.37%, 0.49%, −0.91%, 5.2%, −6.04%, 0.32%, 0.69%, 1.15%, −51.21% and −0.51%.[12] 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 Italy, 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. Again, 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 appears to be no difference in volatility across business cycle phases.

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

$$F^{-1}(\Pr(S_t = 1)) = \tau_0 + \tau_1 D_{99} + \tau_2 y_{\mathrm{GF}} + \tau_3 D_{99} \cdot y_{\mathrm{GF}} + \tau_4 y_{\mathrm{US}} + \tau_5 D_{99} \cdot y_{\mathrm{US}} + \varepsilon_t.$$

Some may suggest that we should include these five variables in the original FTP equation. If we do that, then 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

TABLE 3
ESTIMATES OF FTP MARKOV SWITCHING PARAMETERS

	Eurozone members								Non-eurozone members		
Estimates	Austria	Belgium	Netherlands	Portugal	Ireland	Italy	Greece	Spain	Denmark	Sweden	Switzerland
α_0	3.608	1.573	2.699	1.107	3.969	0.926	0.306	0.992	3.774	2.586	5.668
	(0.015)	(0.030)	(0.002)	(0.232)	(0.022)	(0.112)	(0.717)	(0.023)	(0.000)	(0.003)	(0.001)
α_1	−2.626	−4.154	−2.046	−1.507	−1.480	−4.710	0.232	−0.705	−2.420	−52.365	−6.236
	(0.337)	(0.513)	(0.470)	(0.568)	(0.654)	(0.398)	(0.963)	(0.395)	(0.187)	(0.008)	(0.050)
ϕ_0	−0.770	−1.088	−1.340	−0.439	−0.479	−0.627	−1.135	−0.473	−1.183	−0.972	−1.123
	(0.002)	(0.000)	(0.000)	(0.007)	(0.006)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
γ_1	−0.078	0.069	0.125	−0.472	−0.396	−0.300	−0.017	−0.082	−0.171	−0.090	0.032
	(0.726)	(0.504)	(0.484)	(0.003)	(0.003)	(0.030)	(0.891)	(0.447)	(0.000)	(0.296)	(0.865)
γ_2	0.124	0.072	−0.006	−0.262	−0.206	−0.193	−0.058	0.227	−0.271	−0.040	−0.018
	(0.369)	(0.346)	(0.959)	(0.104)	(0.018)	(0.061)	(0.483)	(0.009)	(0.000)	(0.506)	(0.872)
h_0	17.927	37.887	31.187	23.458	37.333	25.050	35.631	2.669	0.010	57.194	31.238
	(0.017)	(0.000)	(0.005)	(0.097)	(0.004)	(0.000)	(0.001)	(0.005)	(0.112)	(0.000)	(0.001)
h_1	89.122	401.166	141.424	124.579	332.468	97.722	267.194	41.587	323.356	272.011	158.586
	(0.002)	(0.093)	(0.036)	(0.071)	(0.000)	(0.168)	(0.076)	(0.000)	(0.000)	(0.565)	(0.000)
$\tilde{p}$	0.855	0.643	0.906	0.754	0.881	0.911	0.456	0.982	0.978	0.288	0.869
	(0.296)	(0.317)	(0.436)	(0.290)	(0.123)	(0.666)	(0.009)	(0.400)	(0.209)	(0.011)	(0.106)
$\tilde{q}$	0.897	0.958	0.956	0.868	0.884	0.988	0.891	0.919	0.634	0.975	0.885
	(0.116)	(0.296)	(0.288)	(0.197)	(0.084)	(0.465)	(0.295)	(0.286)	(0.091)	(0.318)	(0.053)
Log-likelihood	−343.22	−359.74	−354.90	−353.21	−406.37	−307.15	−362.98	−320.82	−429.00	−373.26	−380.09

Notes
Numbers in parentheses are *p*-values.

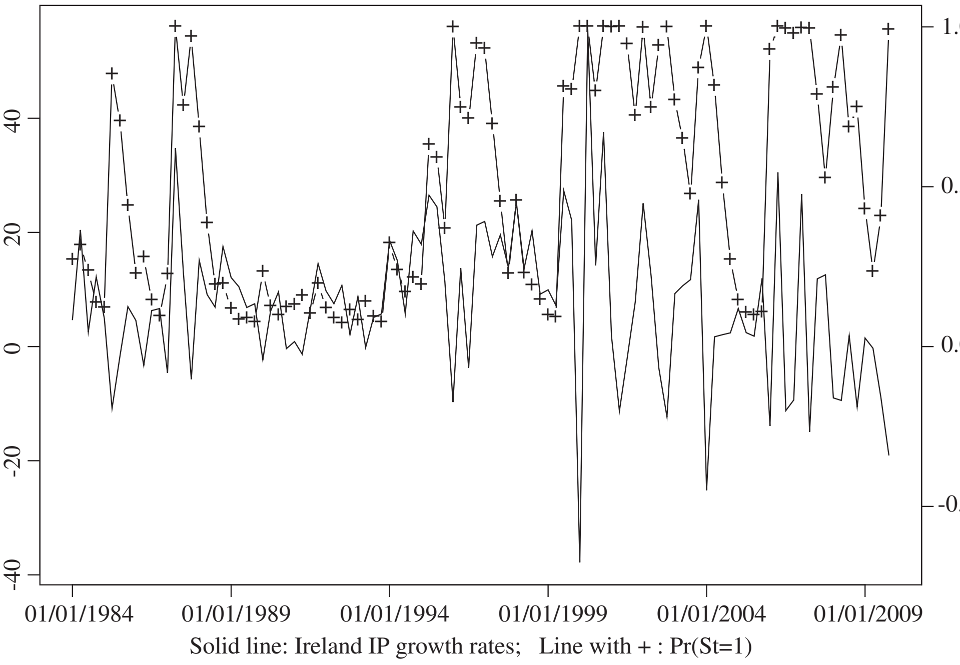

FIGURE 1. Ireland IP growth rates and Pr $(S_t = 1)$

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/expansion. As the FTP MLE estimates are obtained, we are able to calculate the probability of $S_t = 1$ or $S_t = 0$

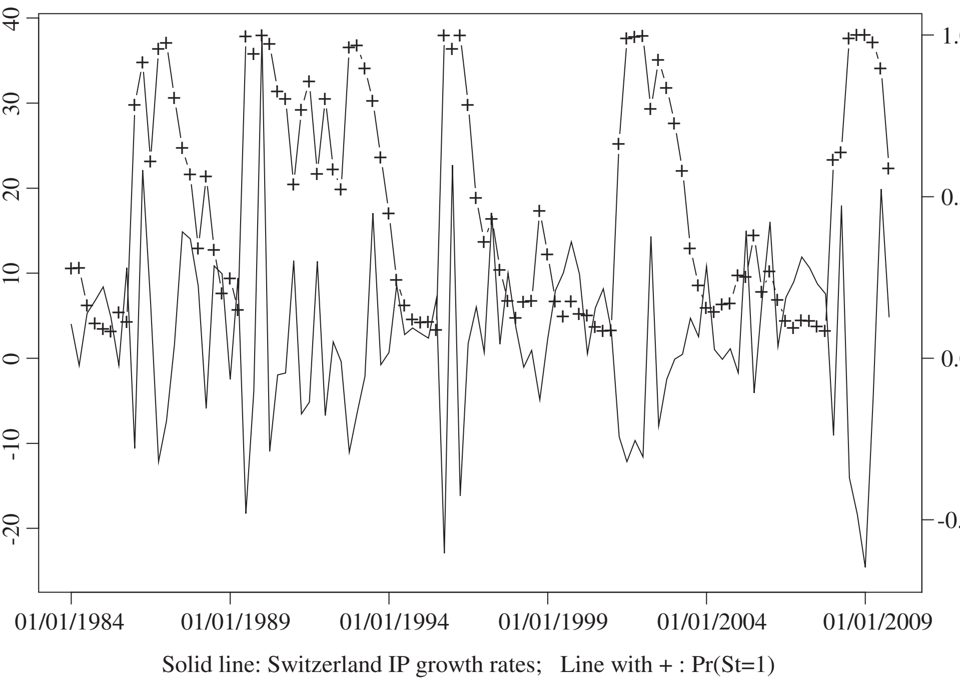

FIGURE 2. Switzerland IP growth rates and Pr $(S_t = 1)$

(i.e. being in a recession or an expansion) for each time period.[14] Thus instead of incorporating all these variables in the original 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 IP growth rates and the probability of being in recession ($S_t = 1$) for Ireland and Switzerland. The left-hand vertical axis denotes the IP growth rate, while the right-hand vertical axis indicates the probability.

The results of the probit model are shown in Table 4. As displayed, for the Netherlands, Greece and Denmark, the null of no US impact could not be rejected at the 5% significance level, thus the two US variables were not included in the probit estimation. Belgium, Ireland, Italy and Spain, in contrast, all appear to have become more synchronized with the USA since the euro's creation (estimates of τ_5 are negative and significant for these countries). Of course, our key parameter of interest is τ_3. The estimate of τ_3 is insignificant for Austria, Belgium, Portugal, Italy and Spain, indicating that joining the euro has led to no change in BCS with the centre for these five nations. However, the estimate is positive and significant for Ireland, and negative and significant for the Netherlands and Greece. These results indicate a *decrease* in BCS for Ireland and Greece, but an increase in BCS for the Netherlands. Again, when the estimate of α_1 is negative, a positive (negative) estimate of τ_3 implies an increase (a decrease) in the probability of being in regime 1 (i.e. a recession regime). For Ireland, the estimate of τ_3 is 0.151, while for the Netherlands the estimate of τ_3 is -0.099. For Greece, the estimate of α_1 is positive, thus regime 1 is an expansion regime; a negative and significant τ_3 implies that after euro adoption, an increase in the centre output has a negative effect on 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 centre of the eurozone. However, for the case of Denmark, the coefficient is negative and significant, so an increase in central eurozone output growth since the creation of the euro now has a significantly larger effect in lowering the probability of a Danish recession.

Once again, joining the euro has led to no increase in BCS with respect to the centre for seven of these eight eurozone nations, and in two cases it led to a decrease in BCS. The Netherlands is the only one that showed an increase in BCS with the centre after euro adoption. Failing to join the euro has led to no loss of BCS, and in one case, BCS with the centre of the eurozone 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. The Netherlands is the one exception, in that its results actually support the idea that joining a monetary union raises BCS—clearly in the FTP model but not so clearly in TVTP model (i.e. a negative but insignificant θ_{p2}). In addition to noting that the Netherlands was an exception, we also emphasize that for both methods, there was no loss of BCS for any of the three non-euro countries. Finally, with both techniques, one non-eurozone member experienced an increase in BCS while retaining its own currency.

TABLE 4
ESTIMATES OF THE PROBIT MODEL

	Eurozone members								Non-eurozone members		
Estimates	Austria	Belgium	Netherlands	Portugal	Ireland	Italy	Greece	Spain	Denmark	Sweden	Switzerland
τ_1	0.820	0.068	0.408	0.038	1.809	0.497	−0.208	0.218	−0.368	−0.947	−0.537
	(0.000)	(0.642)	(0.048)	(0.857)	(0.000)	(0.007)	(0.233)	(0.379)	(0.091)	(0.008)	(0.068)
τ_2	−0.042	−0.006	0.013	−0.032	−0.017	0.003	0.026	−0.039	0.003	0.044	−0.022
	(0.230)	(0.806)	(0.223)	(0.152)	(0.294)	(0.678)	(0.112)	(0.220)	(0.859)	(0.165)	(0.513)
τ_3	−0.003	0.027	−0.099	0.037	0.151	0.041	−0.072	0.015	−0.091	−0.014	−0.004
	(0.952)	(0.429)	(0.000)	(0.306)	(0.002)	(0.314)	(0.008)	(0.819)	(0.001)	(0.793)	(0.952)
τ_4	0.059	0.057		−0.046	0.089	0.040		−0.008		−0.128	−0.093
	(0.025)	(0.097)		(0.111)	(0.000)	(0.000)		(0.686)		(0.011)	(0.007)
τ_5	−0.068	−0.175		−0.061	−0.240	−0.219		−0.128		0.058	−0.065
	(0.373)	(0.013)		(0.137)	(0.007)	(0.002)		(0.030)		(0.536)	(0.306)
τ_0	−0.634	−1.690	−0.897	−0.253	−0.775	−2.048	−0.852	−1.783	−0.489	0.038	0.594
	(0.000)	(0.000)	(0.000)	(0.083)	(0.000)	(0.000)	(0.000)	(0.000)	(0.007)	(0.896)	(0.014)
p-value of H_0:$\tau_4 = \tau_5 = 0$	(0.080)	(0.039)	(0.144)	(0.001)	(0.001)	(0.000)	(0.308)	(0.047)	(0.593)	(0.026)	(0.001)

Notes
Numbers in parentheses are *p*-values.

TABLE 5
IMPACTS OF CENTRAL EURO IP OR US IP ON PR ($S_T = 1$)

			One SD change in both ΔY_{GF} and ΔY_{US}	One SD change in ΔY_{GF} alone	One SD change in ΔY_{US} alone
Eurozone	Austria	Δ Pr ($S_t = 1$) before 1999:Q1	−0.019 (0.791)	−0.096 (0.246)	0.077 (0.037)
		Δ Pr ($S_t = 1$) after 1999:Q1	−0.181 (0.007)	−0.161 (0.172)	−0.021 (0.898)
	Belgium	Δ Pr ($S_t = 1$) before 1999:Q1	0.023 (0.499)	−0.005 (0.803)	0.027 (0.196)
		Δ Pr ($S_t = 1$) after 1999:Q1	−0.050 (0.019)	0.020 (0.452)	−0.072 (0.144)
	Netherlands	Δ Pr ($S_t = 1$) before 1999:Q1		0.022 (0.245)	
		Δ Pr ($S_t = 1$) after 1999:Q1		−0.276 (0.000)	
	Portugal	Δ Pr ($S_t = 1$) before 1999:Q1	−0.142 (0.008)	−0.077 (0.157)	−0.065 (0.073)
		Δ Pr ($S_t = 1$) after 1999:Q1	−0.227 (0.005)	−0.031 (0.680)	−0.197 (0.002)
	Ireland	Δ Pr ($S_t = 1$) before 1999:Q1	0.077 (0.078)	−0.039 (0.311)	0.115 (0.001)
		Δ Pr ($S_t = 1$) after 1999:Q1	0.087 (0.059)	0.297 (0.006)	−0.211 (0.088)
	Italy	Δ Pr ($S_t = 1$) before 1999:Q1	0.011 (0.003)	0.001 (0.677)	0.010 (0.000)
		Δ Pr ($S_t = 1$) after 1999:Q1	−0.073 (0.039)	0.045 (0.376)	−0.123 (0.112)
	Greece	Δ Pr ($S_t = 1$) before 1999:Q1		0.049 (0.159)	
		Δ Pr ($S_t = 1$) after 1999:Q1		−0.095 (0.059)	
	Spain	Δ Pr ($S_t = 1$) before 1999:Q1	−0.026 (0.084)	−0.026 (0.095)	−0.000 (0.993)
		Δ Pr ($S_t = 1$) after 1999:Q1	−0.118 (0.117)	−0.036 (0.436)	−0.078 (0.149)
Non-eurozone	Denmark	Δ Pr ($S_t = 1$) before 1999:Q1		0.008 (0.858)	
		Δ Pr ($S_t = 1$) after 1999:Q1		−0.221 (0.000)	
	Sweden	Δ Pr ($S_t = 1$) before 1999:Q1	−0.052 (0.556)	0.138 (0.086)	−0.189 (0.005)
		Δ Pr ($S_t = 1$) after 1999:Q1	−0.028 (0.497)	0.013 (0.869)	−0.041 (0.691)
	Switzerland	Δ Pr ($S_t = 1$) before 1999:Q1	−0.202 (0.000)	−0.098 (0.131)	−0.105 (0.023)
		Δ Pr ($S_t = 1$) after 1999:Q1	−0.436 (0.000)	−0.107 (0.497)	−0.341 (0.001)

Notes
Numbers in parentheses are p-values.
The impact measured is the effect on the probability of the Markov switching model being in the recession state ($S_t = 1$).

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 eurozone and US 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 1999:Q1, we calculate the change in the probability Pr (S_t = 1) due to a one standard deviation difference in the German–French IP growth and/or the US IP growth. The results can be classified into the following groups. First, for Austria, Portugal, Spain, Sweden and Switzerland, the fourth column of Table 5[15] shows that a one standard deviation increase in both German–French IP growth and US IP growth reduces the probability of being in the recession (and more volatile) regime by 1.9%, 14.2%, 2.6%, 5.2%, 20.2% (respectively) before 1999:Q1, and 18.1%, 22.7%, 11.8%, 2.8%, 43.6% after 1999:Q1. The last two columns show that for Portugal and Switzerland, the reduction of 22.7% and 43.6% after 1999:Q1 is primarily due to the increase of US IP growth rate, not so much by the increase of German–French output growth. Second, for the Netherlands, Greece and Denmark, a one standard deviation increase in German–French IP growth increases the probability of being in regime 1 by 2.2%, 4.9%, 0.8% before 1999:Q1 and decreases the probability of being in regime 1 by 27.6%, 9.5%, 22.1% after 1999:Q1. Note that regime 1 is the contraction regime for the Netherlands and Denmark, while it is an expansion regime for Greece. Thus after euro adoption, the increase of centre output growth indeed significantly decreases the Netherlands' probability of being in a recession regime. Denmark's reaction to the centre output is the same as that of the Netherlands, except that Denmark is a non-euro member. As for Greece, the impact is opposite because regime 1 is an expansion regime for Greece. Third, for Belgium and Italy, after euro adoption, the increase of both German–French and US IP growth decreases the probability of being in a recession regime; and the decrease in probability is primarily due to the US impact. The decomposition shows that the impact of German–French output alone actually increases the probability of being in a recession, even though the estimates are not significant. Finally, for Ireland, a one standard deviation increase in both German–French and US IP growth increases the probability of being in the recession regime both before and after 1999:Q1 by 8% and 9%. But the decomposition in the last two columns of Table 5 shows that after 1999:Q1, US economic expansion actually decreases Ireland's probability of being in recession by 21.1%, but German–French expansion increases that probability by 29.7%.

In terms of the euro's impact on BCS, we examine the impact of German–French growth on the probability of being in regime 1 since the currency's introduction in 1999. For the eight eurozone countries, there is no significant impact in five cases (Austria, Belgium, Italy, Portugal and Spain). There is a significant negative effect for the Netherlands, indicating it being more synchronized with the centre after euro adoption. There is a positive effect, significant at the 5% level for Ireland, and a negative effect, significant at the 10% 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, significant at the 1% level, on Denmark's probability of being in recession, resulting from centre output growth in the post-euro era. Thus, again, joining the euro has led to no impact in BCS with France and Germany for five euro

TABLE 6
SUMMARY OF RESULTS ABOUT EURO ADOPTION ON BCS

	Eurozone members								Non-eurozone members		
Techniques	Austria	Belgium	Netherlands	Portugal	Ireland	Italy	Greece	Spain	Denmark	Sweden	Switzerland
Simple regression	NC	↑ BCS	NC	NC	NC	↑ BCS	NC	NC	NC	↑ BCS	NC
TVTP	NC	NC	NC	↓ BCS	↓ BCS	NC	NC	NC	NC	↑ BCS	NC
FTP/probit	NC	NC	↑ BCS	NC	↓ BCS	NC	↓ BCS	NC	↑ BCS	NC	NC
FTP/Δ Pr ($S_t = 1$)	NC	NC	↑ BCS	NC	↓ BCS	NC	↓ BCS	NC	↑ BCS	NC	NC

Notes
NC indicates that there is no change in BCS with the eurozone's centre economy since euro adoption. ↑ BCS means that the country has an increase in BCS with the centre after the euro adoption, while ↓ BCS indicates that there is a decrease in BCS with the centre.

nations, an increase in one, 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.

In summary, Table 6 provides an overall picture of the results in using different techniques to investigate the endogeneity of the OCA criteria.

IV. Conclusion

The debate over readiness for entry into the eurozone has been split, going back prior to the euro's creation, between those believing that prospective members should fulfil a number of criteria 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 its 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. On the other hand, Frankel and Rose (1996) argue that the act of joining the euro, whatever its other effects, should increase business cycle synchronization among member nations.

The results of this study support the former view more so than the latter. In two cases in each Markov switching model, euro adoption has been followed by less BCS with the core. Moreover, for the three non-euro countries, in two cases there was no loss of BCS associated with retaining their own currencies, while in one case BCS has actually increased since the euro was created, despite (or perhaps because of) failure to join the EMU. In only one case of a euro member—the Netherlands—in only one of our two Markov switching specifications, was there an increase in BCS after joining the EMU.

Results from previous studies shed some possible light on these results. Von Hagen and Neumann (1994) examined readiness for the euro common currency before its creation for a set of eight nations (besides Germany) including the Netherlands, although not Ireland, Greece or Portugal; the authors found, by their metric of real exchange rate variability, that the Netherlands looked 'ready' for euro adoption, compared to some other countries in the sample. Mink *et al.* (2012) examine twelve European nations, including the Netherlands, Ireland, Greece and Portugal, for synchronization over the 1970–2006 period. Their measures show the Netherlands scoring higher on their BCS measures than the other three aforementioned nations.

As Ireland, Greece and Portugal appeared to lose BCS after euro adoption, while the Netherlands seemed to have improved its integration with the core, it may be that in some cases, countries 'ready' for monetary union may experience an increase in BCS by the very act of joining a common currency. In contrast, and as is more often the case, nations that are found to be less ready to join a CU may actually become less integrated with the centre once they join a monetary union.

Despite the turmoil suffered in the eurozone in the wake of the 2008–9 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 likely not increase, and may well decrease cyclical convergence for those countries not already highly integrated. The parallel benefits of 'one-size-fits-all' monetary policy for both mature eurozone members and smaller economies could thus be very questionable.

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NOTES

1. Cited with permission from Professor Rose.
2. As will be shown later, we use five explanatory variables in the probit model. If the five variables are included in the original FTP model, then 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 the FTP/probit approach. In our FTP/probit setting, we assume that variables such as centre 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 centre nations' output growth rates, affect these phenomena.
3. One may argue that the common monetary policy is not suitable for all member countries simultaneously, especially when different members are in different stage of a business cycle. For this very reason, even using moving windows (over time) to calculate the correlation coefficients, these short-term correlation coefficients could not really capture the impacts of the monetary policy and the ensuing converging or diverging output patterns.
4. We 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.
5. Recession periods are thus the periods when $S_t = 1$ and $\alpha_1 < 0$, or when $S_t = 0$ and $\alpha_1 > 0$. For the former case, output growth of the state $S_t = 1$ is lower than the state of $S_t = 0$, because $\alpha_1 < 0$. The same analysis applies to the latter case.
6. 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 (which equals $p - 1$, as shown in the equation above), then 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.

$$G^t = \begin{bmatrix} 0 & 0 & 0 & 0 & 1 & 1 & 1 & 1 \\ 0 & 0 & 1 & 1 & 0 & 0 & 1 & 1 \\ 0 & 1 & 0 & 1 & 0 & 1 & 0 & 1 \end{bmatrix},$$

   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, and for $l = 4$, G is a 32×5 matrix.
7. 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.
8. Regime 1 is a recession regime when α_1 is negative. Later results indicate that most nations have negative estimates of α_1.
9. In the case of TVTP model, 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 $\theta_{p1} = \theta_{p2} = \theta_{q1} = \theta_{q2} = 0$. 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 $\theta_{p1} = \theta_{p2} = \theta_{q1} = \theta_{q2} = 0$ should use the same program with the same number of observations.
10. The joint test result is not really consistent with the individual t-test results, which shows the possibility of multicollinearity issues among these variables. Due to the complexity of the model, we leave it as it is.
11. Note that Belgium has a negative TVTP estimate of α_1, even though the FTP estimate of α_1 is positive. Even so, the Belgium results are consistent in both models: the FTP estimate of h_1 is greater than that of h_0, while the TVTP estimate of h_1 is smaller than that of h_0.
12. The average rate is calculated as $(\alpha_0 + \alpha_1 S_t)/(1 - \rho)$, where $\phi_0 = \rho - 1$.
13. Unlike α_1, h_1 does not measure the differences between regimes 0 and 1.

14. Given

$$\sum_{S_t=0}^{1}\sum_{S_{t-1}=0}^{1}\sum_{S_{t-2}=0}^{1} f(\Delta y_t|S_t, S_{t-1}, S_{t-2}, \Phi_{t-1}) = f(\Delta y_t|\Phi_{t-1}),$$

we can update the probability by

$$f(\Delta y_t|S_t, S_{t-1}, S_{t-2}, \Phi_{t-1})/f(\Delta y_t|\Phi_{t-1}) = f(S_t, S_{t-1}, S_{t-2}|\Delta y_t, \Phi_{t-1}) = f(S_t, S_{t-1}, S_{t-2}|\Phi_t).$$

Then integrating out S_{t-1} and S_{t-2}, we can get Pr (S_t = 1 | Φ_t). For more details, see Kim and Nelson (1999).

15. In Table 5, we report $\hat{F}(\hat{z}_2) - \hat{F}(\hat{z}_1) = \hat{F}(x_2'\hat{\beta}) - \hat{F}(x_1'\hat{\beta})$, 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 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, pp. 1055–6) yields approximate errors.

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