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Stochastic Convergence in the Euro Area

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Abstract

We investigate the dynamics of stochastic convergence of the original Euro Area countries for inflation rates, nominal interest rates, and real interest rates. We test for convergence relative to Germany, taken as the benchmark for core EU standards, using monthly data over the period January 2001 to September 2010. We examine, in a time-series framework, three different profiles of the convergence process: linear convergence, nonlinear convergence, and linear segmented convergence. Our findings both contradict and support convergence. Stochastic convergence implies the rejection of a unit root in the inflation rate, nominal interest rate, and real interest rate differentials. We find that the differentials are consistent with a unit-root hypothesis when the alternative hypothesis is I(0) with a linear trend. But we frequently, but not always, reject the unit-root hypothesis when the alternative is I(0) with a broken trend. We also note that the current financial crisis plays a significant role in dating the breaks.

Journal of Economic Literature Classification: F36, F42, C20, C50

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Introduction

In January 1999, eleven countries of the European Union (EU) formed an economic and monetary union (EMU) that officially became known as the Euro Area or Euro Zone. They abandoned their respective national currencies and relinquished their monetary independence to adopt a common currency, the euro. These countries are Austria, Belgium, Finland, France, Germany, Ireland, Italy, Luxemburg, the Netherlands, Portugal, and Spain. Greece adopted the euro in January 2001. Since 1999, the European Central Bank (ECB) conducts monetary policy. Denmark, Sweden, and the United Kingdom maintain their own currencies and monetary independence.[1]

The EU, as we know it today, was born out of the Maastricht Treaty in 1992, which created the legal, institutional, monetary, and fiscal framework of the EMU (see Eichengreen and Wyplosz, 1993, for more details).[2] The Maastricht Treaty requires that the countries in the Euro Zone achieve real and financial integration.[3] After the inception of the euro, however, we witness a growing number of divergences, especially in connection with the current financial crisis. The relevant literature, however, does not yet determine whether these divergences constitute temporary events or represent the manifestation of more structural phenomena.

This paper analyzes the issue of convergence in the original Euro Area countries, and

[1] Even though these countries lie outside the Euro Area, the static and dynamic effects of the single European currency affect their economies.

[2] Since the Maastricht Treaty, the EU grew with the addition of several countries over time. The EU's first enlargement occurred on May 1, 2004, with the addition of 10 new members: Poland, Hungary, the Czech Republic, Slovakia, Slovenia, Estonia, Latvia, Lithuania, Cyprus, and Malta. The next enlargement round occurred on January 1, 2007, as Bulgaria and Romania joined the EU. The Euro Area also saw increases in its membership. Slovenia became the first of the ten member states, which joined the EU in 2004, to join the Euro Area on January 1, 2007. Then, Cyprus and Malta joined Euro Area on January 1, 2008. Finally, Slovakia became the fourth country to join the Euro Area, since the 2004 enlargement, on January 1, 2008.

[3] In particular, the Maastrich Treaty specified three required steps to establish the EMU. First, by the end of 1993, governments must lift all controls on capital flows within the EU. Second, by 1999, member states preparing to adopt the euro currency must satisfy a set of convergence criteria, which implied coordination of major economic policies across nations. Finally, effective by the beginning of 1999, governments must launch the European Central Bank (ECB), along with the official euro currency for which member-country conversion rates were irrevocably set

assesses the effect of the global financial crisis on the process of convergence.[4] In particular, we consider whether the global financial crisis pulled the twelve economies of the Euro Area together or pushed them apart.

We analyze three aspects of convergence. First, we consider whether inflation differentials persist in the Euro Area.[5] Persistent inflation differences complicate the ECB's monetary policy, as the common interest rate policy may prove too lax for some countries and too strict for others. In such a case, a single monetary policy, "one size fits all," cannot efficiently fight inflation within the Euro Area. Countries with slow growth, high unemployment, and a large trade deficit, such as Greece, Italy, and Portugal, may want to leave the Euro Area to ease monetary conditions and devalue their currencies. Similarly, other countries, such as Germany that prefers tougher monetary policy, may want to leave the Euro Area to pursue a tighter policy. Alternatively, regional inflation differentials may only reflect differences in productivity growth, the well-known Balassa-Samuelson effect. Alberola (2000) provides a full discussion of the potential sources of regional inflation differentials, making a distinction between a benign view of inflation differentials when caused by productivity-related convergence and a more worrying view when inflation differentials result from structural rigidities. In this regard, Duarte and Wolman (2002) show that

[4] Following Schuknecht, von Hagen, and Wolswijk (2010), we distinguish between two phases of the current crisis, a period of market turmoil starting in August 2007 and lasting until August 2008, and the period of the acute crisis starting with the collapse of Lehman Brothers in September 2008.

[5] Empirical evidence on the persistence of inflation differentials appears in Cecchetti, *et al.* (2002) for US cities, Rogers (2001), Berk and Swank (2002), and Ortega (2003) for European countries, Alberola and Marques (1999) and Eijffinger and De Haan (2000) for Spanish provinces. Gregoriou, *et al.* (2007) examine the time-series properties of inflation differentials in twelve EMU countries. A number of papers apply various unit-root and cointegration tests to analyze the persistence of inflation differentials in the Euro Area. Siklos and Wohar (1997) find evidence of convergence for the time period 1974 to 1995. Kocenda and Papell (1997) and Weber and Beck (2005) also report evidence of inflation convergence during the pre-euro period using panel unit-root tests. Busetti, *et al.* (2006) discover convergence during the pre-euro period, but divergence following the introduction of the euro. Rodriguez-Fuentes, *et al.* (2003), contrary to previous studies, identify persistence in inflation differentials in the pre-euro period in eight out of eleven EMU countries. Lopez and Papell (2010), using a new panel-data procedure, uncover strong evidence of convergence among the inflation rates soon after the implementation of the Maastricht treaty and a dramatic decrease in the persistence of the differential after the occurrence of the single currency.

productivity shocks can account for a significant portion of inflation variation across European countries.[6]

Second, we consider whether nominal interest rate differentials also persist in the Euro Area.[7] Since the Euro Area countries share a common currency, exchange-related premium differentials can, by definition, no longer exist in theory, and nominal interest rates on assets with similar characteristics cannot diverge significantly in the Euro Area. Persistence in yield differentials suggests that investors do not regard bonds issued in different Euro Area countries as close substitutes, which, in turn, delivers the financial markets' assessment of the sustainability of the government's fiscal policy, and implies that the real economies of these countries do not follow convergent paths.

Haugh, *et al.* (2009) analyze the recent large movements in the yield spread for sovereign bonds between Germany and other Euro Area countries and conclude that differing fiscal policies, especially their effects on future deficits, and debt service ratios explain significant portions of bond yield spreads in the Euro Area.[8] The important policy making implications of this analysis becomes especially relevant, given the massive expansionary fiscal policies undertaken by Euro Area countries in response to the 2008 recession. The profound deterioration of public finances raises considerable red flags regarding the possible negative effect on interest rates. Particularly, a widespread concern exists in the EMU that the increase in yields on the

[6] A recent ECB (2003) study acknowledges that the monetary policy of the ECB must incorporate the size and persistence of differences in inflation rates, and specifically mentions the possibility that high inflation in some regions could push inflation rates towards deflationary levels elsewhere, which in the presence of downward nominal rigidity may adversely affect economic and eventually even political outcomes.

[7] Several studies provided empirical evidence about nominal interest rate convergence. Karfakis and Moschos (1990), Katsimbris and Miller (1993), Kirchgassner and Wolters (1995), using a bivariate cointegration approach, find no evidence of interest rate convergence. Conversely, Hafer and Kutan (1994), as well as Haug, *et al.*, (2000), using a multivariate cointegration approach, find evidence of partial convergence.

[8] In one ECB (2004) study, Italy, France, and Germany issue 70 percent of the sovereign debt outstanding in the Euro Area, while Spain, Belgium, and the Netherlands issue another 20 percent, and Austria, Finland, Portugal, Greece, Luxemburg, and Ireland issue the remaining 10 percent .

capital markets, driven by the deterioration of national fiscal conditions, will damage the ability to implement the common monetary policy. In early 2010, concerns about fiscal solvency of some Euro Area countries, such as Portugal, Ireland, Greece, and Spain, with the widening of government bond yield spreads and talks of possible bail-outs, disturbed the Euro Area's financial stability. For example, the spread of the Greek 10-year government bond to the 10-year German bond reached a maximum of 329 basis points in February 2010. Similarly, the maximum spread of the Irish 10-year government bond saw 274 basis points in March 2009. Conversely, the spread of the French 10-year government bond peaked in March 2009 at 63 basis points. These differentials reflect the countries' fiscal weaknesses and are especially important in the Euro Area, where member countries can no longer monetize their debt. At the same time, however, the absence of these differentials may only reflect the financial markets' anticipation that the ECB or other member countries will bail out distressed member countries.

Third, we analyze whether real interest rate differentials persist in the Euro Area. Convergence of real interest rates implies that capital movements quickly equalize real rates of returns across the Euro Area, and provides a measure to assess the degree of integration in capital markets.[9]

The concept of convergence inherently relates to economic growth. Therefore, definitions and methodological approaches to convergence come from the empirical growth literature, pioneered by Baumol (1986), Barro (1991), and Barro and Sala-i-Martin (1991, 1992). In contrast to the cross-section notion of convergence, the extant literature also examines a time-

[9] Jenkins and Madzharova (2008), using cointegration techniques, find no evidence of real interest rate convergence, using a sample of 15 EU countries during the period 1999 to 2004. Fountas and Wu (1999) using cointegration tests that allow for endogenously determined regime shift, however, find strong evidence in favor of bilateral real interest rate convergence between several countries in the EU and Germany over 1979 to 1993. Wu and Fountas (2000) also provide strong evidence in favor of bilateral real interest rate convergence between the US and Germany and France. Arghyrou, *et al.* (2009), using time series techniques that allow for endogenously determined structural breaks, report evidence of convergence of real interest rates towards the EMU average.

series notion of convergence. In two seminal contributions, Bernard and Durlauf (1995, 1996), drawing on Carlino and Mills (1993), develop a stochastic approach to convergence, which uses the property of stationarity to evaluate convergence. A stationary series reverts back to its equilibrium value after being disturbed by external shocks. Within this context, the time-series properties of inflation, and nominal and real interest rate differentials are most relevant. A finding that these series are characterized by unit-root behavior indicates stochastic divergence, with shocks affecting the series on a permanent basis.

Bernard and Durlauf (1996) introduce two definitions of convergence – catching up (Definition 1) and equality of long-term forecasts (Definition 2). They demonstrate that countries *i* and *j* do not conform to Definition 2 of convergence when the time-series difference between an economic variable in the two countries contains a non-zero mean or a unit root. That is, absolute stochastic convergence implies that long-run forecasts of the differential between any pair of countries converge to zero, as the forecast interval increases (Oxley and Greasley, 1997). Absolute stochastic convergence exists only if the difference between two time-series is a zero mean stationary process. This can only occur if the paths of the two series are coincident. From this viewpoint, the presence of a deterministic trend or a drift impedes the realization of long-run stochastic convergence.

Second, catching up or conditional convergence, which automatically holds if we can establish the equality of long-term forecasts, focuses on identifying movement toward complete convergence, although it is not yet achieved. Bernard and Durlauf (1996) state that this definition of convergence is violated when history matters or when differences persist into the future. In other words, catching up implies that the differential narrows, but do not necessarily disappear, as the series move through time. This definition also requires that the difference between two

time-series is stationary, but does not preclude the presence of a non-zero mean or a deterministic time trend (Carlino and Mills, 1993).[10] Since the concept of convergence underlying the Maastricht Treaty reflects a gradual process of real and financial integration, we use the second characterization of stochastic convergence (i.e., conditional convergence) and model the bilateral behavior of the inflation rates, nominal interest rates, and real interest rates of the countries in the Euro Area against Germany as a unit-root problem.[11]

We examine three different representations of the convergence process: linear convergence, nonlinear convergence, and linear segmented convergence. First, we apply the efficient version of linear unit-root tests proposed by Ng and Perron (2001). Second, we allow nonlinearities in the convergence process and apply the Kapetanios, et al (2003) unit-root test. Finally, to permit up to two endogenously determined structural breaks, we apply the Perron and Rodriguez (2003) and Lee and Strazicich (2003) unit-root tests.[12]

Data

We use monthly data on interest rates and consumer prices for the original twelve EMU countries that adopted the euro and, thus, a common monetary policy: Austria, Belgium, Finland, France, Germany, Greece, Ireland, Italy, Luxemburg, the Netherlands, Portugal, and Spain. As noted above, Greece adopted the Euro in 2001, while the remaining countries adopted the euro in 1999. To use a common sample, the raw monthly data cover the period from January 2001 to September 2010, and come from the Eurostat databank of the European Central Bank. The data

[10] Li and Papell (1999) postulate an even stronger definition of convergence, called deterministic convergence. This requires the elimination of both deterministic and stochastic trend, and implies that the variables of interest move in parallel in the long-run.

[11] The ECB and numerous papers analyzing convergence in the EMU before or after the onset of the euro (e.g., Fountas and Wu, 1999; Camarero, *et al.*, 2002; Duarte, 2003). According to ECB (2004), we use Germany as the benchmark. Germany's large share in the long-term Euro Area government bond market as well as its achievement of the lowest harmonized long-term German government bond yield at the start of EMU justifies its use as the benchmark.

[12] Existing endogenous structural-break tests can identify at most two breaks.

are "harmonized" to conform to the Maastricht Treaty. The sample consists of 117 observations for each country.

We compute the inflation rate, p_t, as the percentage change of the monthly Harmonized Consumer Price Index (HCPI) over a twelve-month period, and the nominal interest rate, i_t, is the Harmonized 10-Year Government Bond Yield (HGBY).[13] We construct the real interest rate r_t using the ex post version of the Fisher equation. That is,

$$r_t = i_t - p_t, \tag{1}$$

where i_t is the nominal interest rate at time *t*. We use these three measures to construct the series of bilateral inflation differentials $(p_t^i - p_t^G)$, bilateral nominal interest rate differentials $(i_t^i - i_t^G)$, and bilateral real interest rate differentials $(r_t^i - r_t^G)$, where p_t^G, i_t^G, and r_t^G are, respectively, the benchmark measures of the German inflation rate, nominal interest rate, and real interest rate.

To provide a frame of reference for what follows, Figures 1 to 3 plot the three series of differentials $(p_t^i - p_t^G)$, $(i_t^i - i_t^G)$, and $(r_t^i - r_t^G)$ between the original Euro Area countries and Germany from January 2001 to September 2010. These figures reveal some interesting patterns. Visual inspection suggests that all three differential series probably exhibit nonstationarity, and that structural breaks appear to occur generally after 2008.

The series of inflation differentials follow different patterns in different countries. For Greece, the series of inflation differentials takes on positive values throughout the entire sample, while for the remaining countries inflation differentials typically take on positive values for most, but not all, of the sample. For Ireland, the last three years of the sample show a vastly

[13] The yield equals the monthly average of the secondary market daily yields of government bonds with maturities of close to ten years. Since Luxembourg does not issue long-term debt securities with a residual maturity of close to ten years, we use a basket of long-term bonds, which exhibits an average residual maturity of close to ten years. Since a private credit institution issues the bonds, this indicator is, thus, not fully harmonized.

negative inflation differential, suggesting that the real interest rate differential exceeds the nominal interest rate differential. Discernable spikes notably appear in all series, but not with any time commonality.

Interest rate differentials present a more discernible pattern. All series show a limited amount of variability until 2008, when a significant upward spike occurs in all series, immediately followed by a significant reduction, with four exceptions: Greece, Ireland, Portugal, and Spain. In these four cases, only a short-lived reduction in the nominal interest rate differential series occurs, followed within a few months by a larger upward spike. Thus, it is not surprising that these countries currently experience a financial crisis and receive, or remain potential candidates for receiving, a bailout from the EU and the IMF. In all series, the highest nominal interest rate differentials all occur in 2009.

Real interest rate differentials also follow different patterns in different counties. Greece, Ireland, Portugal, and Spain show a distinctive upward trend in the real interest rate differential. Conversely, Austria, Belgium, France, Luxemburg, and the Netherlands show a spike in real interest rate differentials in 2008-2009, followed immediately by a reversal. The highest real interest rate differentials generally occur from 2007 to 2009, with the exception of Finland, where it occurs in 2004.

Tables 1 to 3 present summary statistics of the series. They include the mean, the standard deviation, the skewness and kurtosis, the maximum and minimum (with the date of occurrence below the value), and the Jarque-Bera statistic. In the case of the mean, numbers in parentheses report the *t*-statistics corresponding to the null of zero mean. In the case of the Jarque-Bera statistic, we report the *p*-values. The latter statistic tests the hypothesis of normality of the series. This test statistic, distributed as a chi-square with 2 degrees of freedom, jointly tests

the difference of the skewness and kurtosis of each series from the normal distribution.

The following points deserve mention. First, Greece experiences the largest average inflation and nominal interest rate differentials, and the lowest average real interest rate differential. Second, the inflation differentials' means significantly differ from zero, with the exception of Finland. Third, all nominal interest rate differentials' means also significantly differ from zero. Conversely, the real interest rate differentials' means for Austria, Finland, France, and Ireland do not significantly differ from zero. Fourth, the Jarque-Bera test does not unambiguously reject normality. In the case of inflation rate differentials, the test rejects normality for Belgium, Greece, Luxembourg, and the Netherlands. For nominal interest rate differentials, the test fails to reject normality only in the case of Luxembourg. In all other series, we report both positive skewness and kurtosis, which provides evidence that the distributions exhibit a long right tail. That is, the differentials more likely occur above the mean than below it and the distributions are leptokurtic, implying that the distributions exhibit more mass in the tails than the normal distribution, a phenomenon known as fat tails. Finally, mixed evidence emerges on the normality of real interest rate differentials. The Jarque-Bera test rejects normality for Greece, Ireland, the Netherlands, and Spain. For France, Italy and the Netherlands, skewness and kurtosis are negative. Finally, although we do not report them, the mean and the standard deviation of the series differ noticeably in most cases before and after 2006, which suggests a change in behavior in the last four years.

Convergence Methodology

As emphasized in the introduction, the existence of stochastic convergence relates to the unit-root hypothesis. Stationarity relates to shock persistence in the sense that for a stationary series, shocks only exert temporary effects. On the other hand, for a non stationary series, shocks

possess permanent effects. Thus, if the series of differentials do not possess a unit root, we conclude that evidence exists to support stochastic convergence. In this context, failure to reject the unit-root hypothesis corresponds to failure to provide evidence of stochastic convergence. In this section, we briefly review the methods employed to test stochastic convergence. In what follows, we consider the series $\{d_t\}$ where d_t denotes nominal interest rate differentials, inflation differentials, and real interest rate differentials at time *t*.

First, we examine the convergence of d_t, using the linear tests MZ_t and MP_T developed by Ng and Perron (2001). The tests are constructed using local-to-unity *GLS* detrended data and possess good size and power properties. Most conventional unit-root tests suffer from at least three problems. First, they exhibit low power when the root of the autoregressive polynomial is close to, but less, than unity (DeJong, et al, 1992). Second, most of the tests suffer severe size distortions, when the *MA* polynomial of the first-differenced series comes with a large negative autoregressive root (Schwert, 1989). Third, size distortions and loss of power associate with the selection of the number of lags (Ng and Perron, 1995). Ng and Perron (2001) address these three issues with the M-class of tests, which is robust to these issues. The MZ_t statistic modifies the Z_t statistic of Phillips and Perron (1988), while the MP_T statistic modifies the P_T statistic of Elliott, *et al.* (1996). These statistics are defined as follows:

$$MZ_t = (\kappa / f_0)^{0.5}(T^{-1}\tilde{d}_T^2 - f_0)/2\kappa \text{ and} \tag{2}$$

$$MP_T = (\bar{a}^2\kappa + (1-\bar{a})T^{-1}\tilde{d}_T^2)/f_0, \tag{3}$$

where $\tilde{d}_t$ is the locally detrended data, $\bar{a}$ = -13.5, f_0 is the residual spectrum at frequency zero, and $\kappa = \sum_{t=2}^{T}(\tilde{d}_{t-1})^2 / T^2$.

Next, we examine nonlinear stationarity using the Kapetanios, *et al.* (2003) unit-root test,

a nonlinear version of the Augmented Dickey-Fuller test, based on an exponentially smooth-transition autoregressive model (ESTAR). The Ng and Perron (2001) tests assume that the adjustment toward equilibrium follows linear dynamics. Many situations exist, however, where nonlinear adjustments may occur. For example, policy actions may differ depending on the direction of the deviations from equilibrium, or may take place only when the deviation from equilibrium exceeds a threshold. Kapetanios, *et al.* (2003) argue that nonlinear data generation processes (DGP) can severely affect the power of linear unit-root tests and may prevent the tests to reject the null hypothesis of unit root. Kapetanios, *et al.* (2003) propose the following univariate ESTAR model:

$$\Delta d_t = \gamma d_{t-1} + \left[1 - \exp(-\theta d_{t-1}^2)\right] + \varepsilon_t . \tag{4}$$

The test focuses on the parameter θ which equals zero under the null and is positive under the alternative. Since γ is not identified under the joint null hypothesis of linearity and a unit-root, testing the null hypothesis of $H_0 : \theta = 0$ against the alternative hypothesis of $H_1 : \theta > 0$ is not feasible. Thus, Kapetanios, *et al.* (2003) reparameterize equation (11) using a Taylor series approximation to obtain:

$$\Delta \hat{d}_t = \beta \hat{d}_{t-1}^3 + \sum_{i=1}^{k} \psi_i \Delta \hat{d}_{t-i} + e_t , \tag{5}$$

where $\hat{d}_t$ is the *OLS* detrended series d_t and *k* is the number of lags used to correct for serially correlated errors. Equation (5) tests the null of a unit root against the alternative of a nonlinear, but globally stationary, process using the *OLS t*-statistic corresponding to β.

Finally, we explicitly address the issue of structural breaks. The unit-root literature (see Perron, 1989; Zivot and Andrews, 1992; Perron, 1997) shows that conventional unit-root tests suffer from a loss of power in the presence of structural breaks and may fail to reject the null

hypothesis of a unit root when the series is stationary around a segmented trend function[14]. We employ tests that allow up to two structural breaks, which are determined "endogenously" from the data, so as to capture the possible changes in the convergence path. Not only do important statistical reasons exist for structural breaks in the data generation process of the differential series, but also important economic reasons exist as well. The financial markets of the Euro Area countries experienced significant structural changes from January 2001 to September 2010, where the global financial and economic crisis, when the subprime crisis hit global financial markets in the third week of July 2007, provided the biggest effect. Until then, analysts viewed the stability and convergence of nominal and real interest rates and inflation differentials as the hallmark of successful financial integration within the Euro Area.

We employ several testing procedures for unit roots that permit structural breaks in the data generating process (*DGP*). The Perron and Rodríguez (2003) test extends the *GLS* detrending approach of Elliott, *et al.* (1996) and Ng and Perron (2001) in the context of one unknown structural change. We also consider the possible existence of two breaks under the alternative hypothesis of stationarity. Lumsdaine and Papell (1997) provide a frequently used test based on the *ADF* statistic. *ADF*-type tests suffer a drawback that the presence of structural breaks under the null affects their size properties. Furthermore, these statistics experience divergence under the null hypothesis. To overcome this issue, we use a two-break minimum-LM statistic proposed by Lee and Strazicich (2003), which is related to the one-break LM unit-root test developed by Amsler and Lee (1995). Unlike conventional unit-root statistics, the minimum-LM statistic does not suffer from bias and spurious rejection in the presence of structural breaks

[14] The seminal work of Perron (1989) demonstrates that the unit- root and structural-break issues are intertwined. Perron shows how the presence of a structural break at a known break date biases standard unit-root tests toward non rejection of the null hypothesis of a unit root. Researchers now call this the "Perron phenomenon."

under null.

Perron and Rodriguez (2003) extend the tests for a unit root developed by Ng and Perron (2001) to consider an endogenous change in the trend function. The Ng and Perron (2001) test modifies a version of the Phillips and Perron (1988) test that corrects the size distortions (as suggested by Perron and Ng, 1996) and improves the power (as suggested by Elliott, *et al.*, 1996). The Perron Rodriguez test is based on the following regression:

$$\Delta\tilde{d}_t = \beta\tilde{d}_{t-1} + \sum_{i=1}^{k}\psi_i\Delta\tilde{y}_{t-i} + \nu_{tk}\,, \tag{6}$$

where $\tilde{d}_t$ represents the detrended series, which come from the local-to-unity GLS detrending and *k* is selected using the Modified Akaike Information Criterion (MAIC). Under the unit-root hypothesis, β=0. The test statistics employed is the $MZ_t^{GLS}(\delta)$ statistic, defined as follows:

$$MZ_t^{GLS}(\delta) = \left(T^{-1}\tilde{d}_T^2 - s^2\right)\left(4s^2T^{-2}\sum_{t=1}^{T}\tilde{d}_{t-1}^2\right)^{-1/2}, \tag{7}$$

where $\delta = T_B/T\,, \delta \in (0,1)$ is the relative position of the timing of the break, and *s* is an autoregressive spectral density estimator of the long-term variance. We estimate the break point by minimizing $MZ_t^{GLS}(\delta)$.

Lee and Strazicich (2003) develop a version of the LM unit-root test to accommodate two structural breaks. The two-break test reduces the potential loss of power to reject the unit root when two breaks occur instead of one. The test includes the two breaks under both the null and the alternative hypotheses, and the rejection of the null unambiguously implies stationarity. Lee and Strazicich (2003) analyze two alternative models, which they call Model A and Model C.

We consider Model C,[15] which allows for two shifts in the level and the trend of d_t so that $Z_t = [1,\ t,\ D_{1t},\ D_{2t},\ DT_{1t},\ DT_{2t}]'$. D_{jt} and DT_{jt}, $j = 1,\ 2$, are dummy variables that denote mean and trend shifts, respectively. $D_{jt} = 1$ for $t \geq T_{Bj} + 1,\ j = 1,\ 2$ and 0 otherwise, and $DT_{jt} = t - T_{Bj}$ for $t \geq T_{Bj} + 1,\ j = 1,\ 2$, and 0 otherwise. T_B represents the time period when a break occurs. In Model C, equations (8) and (9) give the null and alternative hypotheses as follows:

$$d_t = \mu_0 + c_1 B_{1t} + c_2 B_{2t} + e_1 D_{1t} + e_2 D_{2t} + d_{t-1} + \upsilon_{1t}, \text{ and} \tag{8}$$

$$d_t = \mu_1 + \gamma t + c_1 B_{1t} + c_2 B_{2t} + e_1 DT_{1t} + e_2 DT_{2t} + \upsilon_{1t}, \tag{9}$$

where the error terms υ_{1t} and υ_{1t} are stationary processes, and $B_{jt} = 1$ for $t = T_{bj} + 1$, $j = 1, 2$, and zero otherwise. The test assumes the following *DGP*:

$$d_t = \delta' Z_t + \eta_t,\ \eta_t = \beta \eta_{t-1} + \varepsilon_t, \tag{10}$$

where Z_t is a vector of exogenous variables and $\varepsilon_t \sim iid\ N(0, \sigma^2)$. The LM unit-root test statistic comes from the following regression:

$$\Delta d_t = \delta' \Delta Z_t + \phi \tilde{S}_{t-1} + \sum_{i=1}^{k} \psi_i \Delta \tilde{S}_{t-i} + \varepsilon_t, \tag{11}$$

where the detrended series $\tilde{S}_t$ is defined as follows: $\tilde{S}_t = d_t - \tilde{\psi}_x - Z_t \tilde{\delta}$, $t = 2,\ ...,\ T$; $\tilde{\delta}$ equals the coefficient vector in the regression of Δd_t onto ΔZ_t; $\tilde{\psi}_x$ equals $d_1 - Z_1 \tilde{\delta}$, where d_1 and Z_1 correspond to the first observations of d_t and Z_t, respectively. We include the lags $\Delta \tilde{S}_{t-i}$ to correct for serial correlation. Under the null hypothesis of a unit root, $\phi = 0$, and under the alternative, $\phi < 0$. The LM test statistic equals the *t*-statistic $\tilde{\tau}$ for testing the null hypothesis

[15] Sen (2003) provides Monte Carlo evidence that Model C yields more reliable estimates of the break dates than Model A, when the break data are treated as unknown.

$\phi = 0$. We find the two breaks $\lambda_j = T_{Bj}/T,\ j = 1,\ 2$ by a grid search $LM(\tau) = \underset{\lambda}{\text{Inf}}\,\tilde{\tau}(\lambda)$ that minimizes $\tilde{\tau}$.

We complement the univariate analysis with a panel unit-root approach, which significantly increases the power of the test against the null. We first apply two common panel unit-root tests. The first test, suggested by Levin, *et al.* (2002), uses the following model:

$$\Delta d_{it} = \rho_i d_{it-1} + z'_{it}\gamma + u_{it},\ i = 1,...,N;\ t = 1,\ ...,\ T, \tag{12}$$

where z_{it} is the deterministic component and u_{it} is a stationary process. The test assumes that $\rho_i = \rho$ for all *i* and the panel statistics is a t- statistic calculated under the null as follows:

$$t_\rho = \frac{(\hat{\rho} - 1)\sqrt{\sum_{i=1}^{N}\sum_{t=1}^{T}\tilde{d}_{it-1}^2}}{s_u}, \tag{13}$$

where $s_u^2 = \frac{1}{NT}\sum_{i=1}^{N}\sum_{t=1}^{T}\tilde{u}_{it}^2$, and $\tilde{d}_{it}$ and $\tilde{u}_{it}$ are, respectively, d_{it} and u_{it} corrected for the deterministic component.

The second test, developed by Im, *et al.* (2003), allows ρ_i to vary across the different panel units and uses the average of the individual unit-root test statistics obtained from Augmented Dickey-Fuller regressions. That is,

$$\bar{t} = \frac{1}{N}\sum_{i=1}^{N} t_{\rho i}, \tag{14}$$

where $t_{\rho i}$ is the individual t-statistic for testing the unit-root hypothesis. Im, *et al.* (2003) show that $\bar{t}$ is normally distributed under the null hypothesis, and use estimates of its mean and variance to convert $\bar{t}$ into a standard normal statistic so that conventional critical values can evaluate its significance. The two panel tests also differ in how they treat the alternative

hypothesis. In Levin, *et al.* (2002), the alternative hypothesis assumes that all of the series are trend stationary, while in Im, *et al.* (2003), the alternative hypothesis assumes that at least one, rather than all, of the series are trend stationary.

Finally, we apply the panel LM unit-root test recently proposed by Im, *et al.* (2005). In contrast to conventional panel unit-root tests, the distribution of the panel LM test statistic proves invariant to break-point nuisance parameters. As such, the critical values that apply to the panel LM unit-root test without breaks are also valid for the panel LM unit-root test with breaks, regardless of the location and number of breaks in each cross-section. The panel LM test statistic comes by averaging the optimal univariate LM test statistics estimated for each series. That is, define

$$\overline{LM}_{NT} = \frac{1}{N}\sum_{i=1}^{N} LM_i \tag{15}$$

to equal the mean of the individual LM statistics obtained from equation (11), where N is the number of series. Im, *et al.* (2005) compute the following standardized panel LM test statistic:

$$\Gamma_{LM} = \frac{\sqrt{N}\left(\overline{LM}_{NT} - E\left(L_T\right)\right)}{\sqrt{V\left(L_T\right)}}, \tag{16}$$

where $E(L_T)$ and $V(L_T)$ denote the expected value and the variance of LM_i, respectively, under the null hypothesis that all ϕ are zero. The numerical values for $E(L_T)$ and $V(L_T)$ appear in Im, *et al.* (2005, Table 1). The asymptotic distribution of Γ_{LM} depends only on N and T, and follows the standard normal distribution.

Empirical Results

Table 4 displays the results of linear unit-root tests without structural breaks. We select the lag order using the modified Akaike information criteria (MAIC), assuming an upper bound of $k^* =$

12. Using the 5-percent significance level, we find poor evidence for convergence, since the tests cannot reject the presence of a unit root in favor of the trend stationary alternative. We reject nominal and real interest rate convergence in 10 out of 11 cases (Austria is the only exception), while we reject inflation convergence in all 11 cases. For Greece, the test results suggest the possibility that interest rate differentials follow an explosive process.

Well-known power gains exist in unit-root tests using panel techniques. Our panel results, however, do not differ from the univariate estimates. The Levin, *et al*. (2002, *LLC*) and Im, *et al.* (2003, *IPS*) panel unit-root tests yield test statistics that cannot reject the unit-root hypothesis. For the interest rate differentials, we cannot reject the null hypothesis of a unit root using the *LLC* and *IPS* test statistics (See Table 4). When we exclude Greece, Ireland, Portugal, and Spain, the panel tests still cannot reject the unit-root hypothesis -- the *LLC* test statistic is 0.727 (*p*-value = 0.766) and the *IPS* test statistic is 0.518 (*p*-value = 0.697). Using only Greece, Ireland, Portugal, and Spain, the panel unit-root tests still do not reject the null hypothesis -- the *LLC* test statistic is 10.476 (*p*-value = 1.000) and the *IPS* test statistic is 6.200 (*p*-value = 1.000).

For inflation differentials, we also cannot reject the null hypothesis of a unit root using the *LLC* and *IPS* test statistics (See Table 4). When we exclude Greece, Ireland, Portugal, and Spain, the panel unit-root tests do not reject the null hypothesis -- the *LLC* test statistic is 0.963 (*p*-value = 0.832) and the *IPS* test statistic is -0.417 (*p*-value = 0.338). Using only Greece, Ireland, Portugal, and Spain, the panel unit-root tests cannot reject the null hypothesis -- the *LLC* test statistic is 2.365 (*p*-value = 0.991) and the *IPS* test statistic is 1.257 (*p*-value = 0.895).

Finally, for the real interest rate differentials, we also cannot reject the null hypothesis of a unit root using the *LLC* and *IPS* test statistics (See Table 4). When we exclude Greece, Ireland, Portugal, and Spain, the panel unit-root tests still do not reject the null hypothesis -- the *LLC* test

statistic is 1.124 (p-value = 0.869) and the *IPS* test statistic is -1.066 (p-value = 0.143). Using only Greece, Ireland, Portugal, and Spain, the panel unit-root tests, once again, cannot reject the null hypothesis -- the *LLC* test statistic is 2.684 (p-value = 0.996) and the *IPS* test statistic is 1.585 (p-value = 0.943).

We note that the panel unit-root tests improve the power of the tests. Moreover, since both tests cannot reject the null, these rejections call into question our limited finding of convergence for the nominal and real interest rate differentials in Austria and inflation rate differentials in Luxembourg.[16] In addition, we uncover the same findings when we exclude those countries identified in the popular press as experiencing significant issues with their sovereign debt -- Greece, Ireland, Portugal, and Spain or when consider these countries by themselves.

Table 5 reports the results of the Kapetanios, *et al.* (2003) nonlinear unit-root test, with the number of lags k determined using the Akaike information criterion (AIC), assuming an upper bound of $k^* = 12$. When accounting for nonlinearity, the test adds three more rejections of the null hypothesis of unit root at the 5-percent level, while losing the two cases for Austria identified by the linear unit-root tests in Table 4. We find nonlinear convergence of interest rate differentials for Belgium and Finland, and of real interest rate differentials for Greece. The nonlinear tests provide a marginal increase in support for convergence when compared to the linear tests.

Table 6 reports the results of using the two-break LM unit-root test. In all cases, the maximum number of lags is $k^* = 12$. We reject interest rate convergence in only 1 of the 11 series using a 5-percent level of significance. In other words, we cannot reject the unit-root null

[16] Remember that the LLC alternative hypothesis is that all panel units exhibit stationary series whereas the IPS alternative hypothesis only requires that one unit exhibit stationary series. See the discussions in the previous section.

in only one country -- Luxemburg. We also reject inflation rate convergence at the 5-percent level in 7 of 11 countries. That is, we cannot reject the unit-root hypothesis for 7 countries -- Austria, Belgium, Finland, Ireland, Italy, Portugal, and Spain. For 4 of these countries, Finland, Ireland, Portugal, and Spain, we can accept convergence by rejecting the unit-root hypothesis, but only at the 10 percent significance level. We can accept real interest rate convergence in only two cases, France and the Netherlands. For interest rate differentials, the most frequently identified period of breaks occurs in the years 2008 to 2009. Similarly, in case of inflation differentials, the highest frequency occurs in 2008, while for real interest rate differentials the breaks most frequently occur in 2006 to 2008.

The inclusion of endogenously determined structural breaks in the data generating process leads to slightly different findings. Now, we generally find convergence for nominal interest rates, but limited evidence of convergence for inflation rate and real interest rate differentials, employing the 5-percent significance level. When we relax the significance level to 10 percent, we find much more evidence of convergence of inflation rate differentials (i.e., 8 rather than 4 countries) as well as more evidence of real interest rate differential convergence (i.e., 4 rather than 2 countries).

Further examination of these results reveals that the two structural breaks do not prove significant at the 5-percent level in all cases. The inclusion of two breaks when only one actually occurs affects the power to reject the null. For interest rate differentials, the one-break test is appropriate for 5 countries -- Austria, France, Greece, Ireland, and Luxemburg. For inflation differentials, one structural break is appropriate for four countries – Belgium, France, Portugal, Spain, and for real interest rate differentials, it is appropriate only for Greece.

Tables 7 and 8 present the one-break unit-root tests. While we focus on the series that

failed to satisfy the two-break test, we report the test results for all series. Table 7 reports the results of the Perron and Rodriguez (2003) unit-root test based on Model II, while Table 8 reports the Lee and Strazicich (2004) results based on Model C. Both test results support interest rate convergence for Austria and Ireland, but not for France and Luxemburg. They do not identify, however, the break dates at the same time, since for Austria the Lee and Strazicich test dates the break in 2008 and the Perron and Rodriguez test dates the break in 2005.[17] Greece provides mixed evidence with the Perron-Rodriguez supporting convergence and the Lee-Strazicich not supporting convergence. Furthermore, in no instances do the one-break Perron-Rodriguez or Lee-Strazicich results support inflation convergence for Belgium, France, Portugal, and Spain or real interest rate convergence for Greece.

Combining the results from Tables 6, 7 and 8 give the following conclusions. While Table 6 identifies only the nominal interest rate differential of Luxembourg as exhibiting non-stationarity, the consideration of the Perron-Rodriguez and Lee-Strazicich one-break unit-root tests adds France to this non-stationary list while the Lee-Strazicich test alone suggests that Greece exhibits a non-stationary nominal interest rate differential. For the tests of inflation rate differentials, the Perron-Rodriguez and Lee-Strazicich one-break unit-root tests remove France, Portugal and Spain from the list of stationary series.[18] Finally, the Perron-Rodriguez and Lee-Strazicich one-break unit-root tests do not alter the findings for real interest rate differentials in Table 6,

The panel LM results provide further significant evidence of convergence (See Table 6). The Γ_{LM} statistics rejects the null hypothesis of a unit root for interest rate differentials, inflation

[17] No a priori reason exists to expect the break dates estimated using the Lee-Strazicich and Perron-Rodriguez procedures to coincide.

[18] Table 6 reports that the inflation rate differential in France rejects the unit root hypothesis at the 5-percent level whereas this differential for both Portugal and Spain reject the unit-root hypothesis at the 10-percent level.

rate differentials, and real interest rate differentials. Dichotomizing the sample does not affect the results. When we exclude Greece, Ireland, Portugal, and Spain, we also can reject the null hypothesis -- the Γ_{LM} statistic is -18.774 for nominal interest rate differentials, -16.2652 for inflation differentials and -13.774 for real interest rate differentials. Using only Greece, Ireland, Portugal and Spain, we still can reject the null hypothesis -- the Γ_{LM} statistics is -25.006 for nominal interest rate differentials, -13.381 for inflation differentials, and -10.601 for real interest rate differentials. This panel unit-root test that incorporates different structural breaks uses the alternative hypothesis that at least one panel unit exhibits stationary series. Given that we find at least one country for each differential series that exhibits convergence, the panel unit-root tests confirm the country-by-country findings.

Conclusions

We investigate whether the time series of the inflation rate, the nominal interest rate, and the real interest rate differentials (with respect to Germany, taken as the benchmark to represent core EU standards) in the original Euro Area countries converge stochastically. The methodological framework builds on the literature on growth convergence and brings together several econometric techniques to address the unit-root properties of these variables. In particular, we employ three modeling paradigms. The first paradigm relies on the linear unit-root tests that do not permit structural breaks to exist in the deterministic component of the trend function. The second relies on unit-root tests that allow for nonlinear adjustment. The third relies on unit-root tests that endogenously determine structural breaks. The use of the structural break approach in this context is novel and provides results that generate new insights into the convergence process.

The study covers the period January 2001 to September 2010. This data set extends the

sample used in previous empirical analyses and at the same time, highlights the effects of the current crisis. We find significant evidence of structural breaks using the two break LM test of Lee and Strazicich (2003). If less than two breaks are significant, we use the one break unit-root tests of Perron and Rodriguez (2003) and Lee and Strazicich (2004). The findings of the univariate one- and two-break unit-root tests overturn the results of the linear and nonlinear tests that do not allow for structural breaks, whereas the results of the Im, *et al.* (2005) panel unit-root test that allow for structural breaks overturn the findings of Levin, *et al*. (2002) as well as Im, *et al*. (2003) that do not allow for structural breaks.

Given that we find strong evidence of structural breaks in the inflation, nominal interest, and real interest rate differential series, we cannot place much confidence in our linear unit root tests without structural breaks. Focusing on the findings that Tables 6, 7, and 8 report, we draw the following conclusions at the 5-percent significance level. We find that for the real interest rate differential, only two countries exhibit convergence – France and the Netherlands. For the inflation rate differentials, three countries show convergence – Greece, Luxembourg, and the Netherlands. Finally, nine countries display convergence for the nominal interest rate differentials. France and Luxembourg provide the exceptions. In sum, we discover little evidence of convergence for inflation rate and real interest rate differentials, but much evidence of convergence for nominal interest rate differentials.

What conclusions can we draw from our findings? The primary objective of the ECB has been to achieve and maintain price stability. Previous empirical evidence (e.g., Gali 2003, Rodriguez-Fuentes and Olivera-Herrera 2003, Angeloni and Dedola 1999, and Batini 2002, among others) have shown that this objective was fulfilled in the 1990s. The new empirical evidence presented in this paper reveals mixed support for the hypothesis that inflation and

interest rate differentials have stochastically converged in the period January 2001-March 2010. The implementation of a single monetary policy did not eliminate these differentials in all countries. This finding, however, does not break new ground, as researches find persistence of inflation differentials in both the Spanish regions and the US (Rodriguez-Fuentes, *et al.*, 2004).

Our findings, however, highlight the immediacy of the current debate in the Euro Area between Euro-supporters and Euro-skeptics as Greece, Ireland, Portugal, and Spain face mounting sovereign debt problems. Although the European Central Bank implements monetary policy, individual countries remain sovereign over fiscal policy. And therein lies the problem. Unlike most states in the US, individual countries in the Euro Area can run fiscal deficits to stimulate their local economies. As a result, the policies of the Southern tier countries (Greece, Ireland, Italy, Portugal, and Spain) caused them to become less competitive with the Northern tier countries (Germany, France, the Netherlands, and Belgium) over time. Tied to a common currency, the Southern tier countries cannot devalue their currencies to regain their competitive position. Thus, either fiscal austerity and lower prices in the Southern tier or fiscal expansion and higher prices in the Northern tier provide the basic options to rebalance the competitive positions of the Southern tier countries. Of course, the Northern tier countries do not want to inflate their economies, leaving it to the Southern tier countries to contemplate austerity and recession. The Euro-skeptics see the potential breakup of the Euro Area and a withdrawal of the weaker Southern tier countries from the Euro.

Do these findings imply that the ECB failed to achieve its goal? Do they imply that restrictive monetary policy cannot alone support convergence and that convergence also requires fiscal restraint? Or do they imply that these differentials reflect diverging degrees of nominal rigidities observed across the countries of the Euro Area? Many factors may explain such

rigidities, such as wage dynamics not linked to productivity developments, structural inefficiencies and market rigidities as well as misaligned fiscal policies. While the Northern tier countries possess broadly similar economic structures, the Southern tier present substantial structural differences. Obviously, the solution of these problems does not fall within the domain of monetary policy. Our findings that shocks permanently affect inflation and real interest rate differentials suggest that fiscal policy measures may prove significant. National fiscal policies can react to shocks in such a way as to counteract the emergence of differentials. From this perspective, a sound control of public spending may prove an essential element of economic stability. Substantial research on these issues is needed.

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Table 1: Descriptive Statistics: Bilateral Nominal Interest Rate Differentials

Country	Mean	Std. Dev.	Skewness	Kurtosis	Minimum	Maximum	JB Statistic
Austria	0.229*	0.240	1.750	3.073	-0.203956	0.990	105.767
	(10.345)				Mar-10	Mar-09	[0.000]
Belgium	0.266*	0.245	1.489	1.855	0.030	1.110	60.011
	(11.767)				Jan-05	Feb-09	[0.000]
Finland	0.174*	0.178	1.537	2.877	-0.050	0.800	86.408
	(10.556)				Oct-05	Feb-09	[0.000]
France	0.150*	0.142	1.486	1.347	0.020	0.630	51.884
	(11.429)				Jan-06	Mar-09	[0.000]
Greece	0.933*	1.606	3.388	12.151	0.130	3.290	943.598
	(6.282)				Jan-05	Feb-10	[0.000]
Ireland	0.500*	0.829	1.943	2.931	-0.050	2.740	115.510
	(6.523)				Oct-05	Mar-09	[0.000]
Italy	0.460*	0.381	1.623	1.517	0.140	1.550	62.612
	(13.043)				Sep-03	Nov-09	[0.000]
Luxemburg	-0.120†	0.720	0.010	-1.001	-0.600	1.410	4.885
	(1.797)				Apr-03	Apr-09	[0.087]
Netherlands	0.142*	0.152	1.870	3.580	-0.020	0.690	130.666
	(10.075)				Jul-06	Nov-09	[0.000]
Portugal	0.456*	0.649	3.033	10.090	0.000	1.660	675.620
	(7.599)				Jan-05	Mar-09	[0.000]
Spain	0.296*	0.403	2.340	5.842	0.000	1.100	273.177
	(7.931)				Aug-05	Feb-09	[0.000]

Note: The sample period runs from January 2001 through September 2010. The JB-statistic is the Jarque and Bera (1980) test for normality. Numbers in parentheses are t-statistics, testing the null hypothesis that the mean equals zero. Numbers in brackets are *p*-values. The symbols *, **, and † mean significantly different from zero at the 1-. 5-. and 10-percent levels, respectively.

Table 2: Descriptive Statistics: Bilateral Inflation Rate Differentials

Country	Mean	Std. Dev.	Skewness	Kurtosis	Minimum	Maximum	JB Statistic
Austria	0.231*	0.358	-0.189	-0.204	-0.624	1.003	0.896
	(6.986)				Feb-06	Aug-02	[0.639]
Belgium	0.439*	0.727	0.585	0.846	-1.215	2.500	10.154
	(6.530)				Sep-07	Sep-08	[0.000]
Finland	0.057	1.056	-0.270	-0.599	-2.177	2.255	3.171
	(0.583)				Dec-04	Dec-08	[0.205]
France	0.243*	0.547	-0.024	-0.597	-1.039	1.360	1.751
	(4.817)				Sep-07	Feb-03	[0.417]
Greece	1.748*	0.868	1.095	1.641	0.241	3.053	36.498
	(21.769)				Oct-07	Feb-03	[0.000]
Ireland	0.713*	1.863	-0.422	-0.144	-3.625	3.952	3.580
	(4.142)				Mar-10	May-05	[0.167]
Italy	0.623*	0.642	0.482	-0.172	-0.903	2.166	4.683
	(10.493)				Sep-07	May-03	[0.096]
Luxemburg	1.027*	0.866	-0.505	-0.503	-1.008	2.655	6.205
	(12.830)				Apr-09	Oct-05	[0.045]
Netherlands	0.540*	1.394	0.759	-0.573	-1.557	3.746	12.819
	(4.186)				Nov-07	Dec-01	[0.002]
Portugal	0.842*	1.262	0.177	-0.893	-1.578	3.396	4.499
	(7.215)				Jun-09	Mar-01	[0.105]
Spain	1.193*	0.864	-0.396	-0.054	-0.980	2.868	3.079
	(14.945)				Jun-09	Dec-02	[0.214]

Note: See Table 1. The symbols *, **, and † mean significantly different from zero at the 1-. 5-. and 10-percent levels, respectively.

Table 3: Descriptive Statistics: Bilateral Real Interest Rate Differentials

Country	Mean	Std. Dev.	Skewness	Kurtosis	Minimum	Maximum	JB Statistic
Austria	-0.002	0.384	0.395	0.632	-0.888	1.000	4.991
	(0.063)				Apr-05	Apr-09	[0.083]
Belgium	-0.173*	0.698	-0.082	0.942	-2.062	1.694	4.459
	(2.675)				Jun-08	Jun-09	[0.108]
Finland	0.117	0.934	0.525	-0.377	-1.585	2.277	6.066
	(1.355)				Dec-08	Dec-04	[0.049]
France	-0.093†	0.577	-0.058	-0.559	-1.300	1.179	1.590
	(1.745)				Feb-03	Sep-07	[0.452]
Greece	-0.815*	1.330	1.502	3.185	-2.763	1.981	93.453
	(6.630)				Feb-03	Apr-09	[0.000]
Ireland	-0.213	2.504	0.969	0.184	-3.841	5.161	18.459
	(0.921)				Feb-03	Dec-09	[0.000]
Italy	-0.163**	0.765	-0.623	-0.270	-1.946	1.253	7.918
	(2.301)				May-03	Sep-07	[0.019]
Luxemburg	-1.147*	1.371	0.538	-0.110	-2.894	2.418	5.696
	(9.049)				Jan-03	Apr-09	[0.058]
Netherlands	-0.398*	1.382	-0.799	-0.520	-3.596	1.677	13.753
	(3.113)				Dec-01	Nov-07	[0.001]
Portugal	-0.386*	1.590	0.289	-0.809	-2.976	2.715	4.819
	(2.627)				Mar-01	Apr-09	[0.090]
Spain	-0.897*	1.124	0.686	-0.267	-2.768	1.786	9.511
	(8.640)				Dec-02	Apr-09	[0.009]

Note: See Table 1. The symbols *, **, and † mean significantly different from zero at the 1-, 5-, and 10-percent levels, respectively.

Table 4: Linear Unit-Root Tests with No Structural Break

	Nominal Interest Rate Differentials			Inflation Rate Differentials			Real Interest Rate Differentials		
Country	MZ_t	MP_T	k	MZ_t	MP_T	k	MZ_t	MP_T	k
Austria	-3.028**	5.011**	11	-1.220	25.914	12	-3.147**	4.653**	1
Belgium	-1.587	14.961	11	-0.927	30.564	12	-2.278	8.648	0
Finland	-2.381	8.030	2	-1.478	20.647	12	-1.356	24.482	12
France	-1.511	18.310	11	-1.882	12.718	12	-1.673	16.189	12
Greece	2.175	73.671	12	-1.027	19.554	1	-0.913	17.481	1
Ireland	-2.670†	7.039	3	-1.650	16.323	1	-0.898	26.310	0
Italy	-0.501	35.033	10	-2.205	9.363	1	-2.235	8.989	1
Luxembourg	-1.169	32.921	4	-2.810†	5.791†	0	-2.015	11.197	12
Netherlands	-2.432	7.701	7	-1.360	20.307	0	-1.510	16.872	0
Portugal	-1.930	10.155	6	-2.175	9.118	2	-2.485	7.379	2
Spain	-0.655	23.703	11	-2.205	9.355	0	-1.821	13.071	0

	Nominal Interest Rate Differentials		Inflation Rate Differentials		Real Interest Rate Differentials	
Panel Tests	Test	*p*-value	Test	*p*-value	Test	*p*-value
LLC	6.409	1.000	2.179	0.985	2.619	0.995
IPS	4.131	1.000	0.417	0.662	0.096	0.538

Note: The Ng and Perron (2001) M-class of tests include the MZ_t statistic and the MP_T statistic. The 1%, 5%, and 10% asymptotic critical values for MZ_t are -3.42, -2.91, and -2.62, respectively, and for MP_T are 4.03, 5.48, and 6.67, respectively. The *LLC* and *IPS* are the panel unit-root tests of Levin, *et al.* (2002) and Im, *et al.* (2003), respectively. The symbols *, **, and † mean significantly different from zero at the 1-. 5-. and 10-percent levels, respectively.

Table 5: Kapetanios Nonlinear Unit-Root Test

	Nominal Interest Rate Differential		Inflation Rate Differential		Real Interest Rate Differential	
Country	***t-statistic***	***k***	***t-statistic***	***k***	***t-statistic***	***k***
Austria	-0.531	12	-0.408	12	0.701	12
Belgium	-6.695*	12	-2.051	12	-1.890	12
Finland	-4.000*	1	-0.740	11	-1.117	12
France	-1.465	12	-0.103	11	0.213	11
Greece	2.890	12	-2.185	11	-3.690**	2
Ireland	-2.716	2	-0.905	11	0.089	11
Italy	-0.050	12	-1.264	11	-1.751	11
Luxembourg	-0.396	8	-0.498	11	-0.915	11
Netherlands	-0.962	5	0.451	11	0.268	11
Portugal	-2.920	5	-2.065	12	-1.608	12
Spain	-3.072	12	-0.028	5	-0.375	5

Note: The 1%, 5%, and 10% asymptotic critical values (case 3) are -3.93, -3.40, and -3.13, respectively. The symbols *, **, and † mean significantly different from zero at the 1-. 5-. and 10-percent levels, respectively.

Table 6: Two-Break Minimum LM Unit-Root Tests, January 2001-September 2010

	Nominal Interest Rate Differentials			Inflation Rate Differentials			Real Interest Rate Differentials		
Country	T_{B1} T_{B2}	*Test Statistic*	*k*	T_{B1} T_{B2}	*Test Statistic*	*k*	T_{B1} T_{B2}	*Test Statistic*	*k*
Austria	May-06	-7.256*	12	Nov-04	-4.726	12	Oct-07	-4.261	12
	Sep-08			Mar-06			Oct-08		
Belgium	Mar-07	-6.466*	12	Mar-08	-3.948	12	Nov-02	-3.943	12
	Dec-08			Aug-09			Aug-06		
Finland	May-06	-5.757**	11	Jan-04	-5.689†	12	Jan-04	-5.510†	12
	Aug-08			Aug-08			Dec-08		
France	Apr-06	-6.249**	12	Feb-03	-6.373*	11	Dec-02	-5.906**	11
	Sep-08			Jan-06			Apr-05		
Greece	Jul-07	-9.561*	10	Jan-04	-6.352*	11	Sep-07	-5.167	4
	Jan-09			Oct-08			Mar-09		
Ireland	Oct-07	-9.568*	11	Oct-04	-5.688†	11	Oct-04	-4.862	11
	Feb-09			Oct-08			Oct-08		
Italy	Apr-08	-6.542*	11	Jan-03	-4.865	12	Jan-03	-4.781	12
	Feb-09			Aug-05			Sep-04		
Luxembourg	Apr-04	-4.369	3	Jun-03	-6.424*	11	May-03	-4.314	12
	Jan-09			Jun-06			Apr-06		
Netherlands	Jun-05	-5.943**	11	Feb-04	-6.627*	11	Jan-04	-6.043**	11
	Aug-08			Sep-08			Mar-06		
Portugal	Feb-08	-11.267*	5	May-04	-5.665†	11	May-04	-4.575	12
	Mar-09			Nov-08			Jun-06		
Spain	Oct-07	-6.731*	8	Dec-06	-5.668†	11	Aug-03	-5.483†	8
	Mar-09			Aug-08			Apr-06		

	Nominal Interest Rate Differentials	Inflation Rate Differentials	Real Interest Rate Differentials
Panel Test	*Test Statistic*	*Test Statistic*	*Test Statistic*
Γ_{LM}	-30.054*	-21.040*	-17.385*

Note: Critical values for the LM two-break unit-root test statistic in Model C depend on the location of the break. The 1%, 5%, and 10% critical values for different values of T_{B1}/T and T_{B2}/T come from Table 2 in Lee and Strazicich (2003). The symbols *, **, and † mean significantly different from zero at the 1-. 5-. and 10-percent levels, respectively.

Table 7: Perron Rodriguez One-Break Unit-Root Tests

	Nominal Interest Rate Differentials			Inflation Rate Differentials			Real Interest Rate Differentials		
Country	MZ_t^{GLS}	***k***	***Break Date***	MZ_t^{GLS}	***k***	***Break Date***	MZ_t^{GLS}	***k***	***Break Date***
Austria	-4.468*	5	Aug-05	-3.705**	1	Aug-05	-3.564**	1	Oct-08
Belgium	-2.704	0	Jun-06	-2.476	0	Jun-09	-2.666	0	Mar-09
Finland	-2.985	3	Jun-04	-2.530	0	Jan-04	-2.786	0	Jan-04
France	-3.057	5	Jun-06	-2.451	1	Mar-04	-2.623	1	Mar-04
Greece	-8.596*	2	Jan-08	-2.663	2	Aug-09	-2.848	0	Mar-09
Ireland	-8.216*	3	Jan-09	-2.154	0	Jun-09	-2.431	0	Jan-08
Italy	-3.569**	1	Mar-08	-2.561	1	Apr-04	-3.055	1	Jan-04
Luxembourg	-2.115	4	Dec-04	-3.066	0	Jul-06	-2.500	0	Oct-06
Netherlands	-3.381†	6	Nov-08	-2.913	0	Jul-04	-3.059	0	Dec-04
Portugal	-9.810*	6	Mar-05	-2.803	2	Jun-09	-3.186	2	Mar-06
Spain	-3.103	2	Sep-07	-2.526	0	Apr-03	-2.518	0	Nov-06

Note: The 1%, 5% and 10% asymptotic critical values (Model II) are -3.93, -3.56, and -3.36, respectively, and come from Table 1 in Perron and Rodriguez (2003). The symbols *, **, and † mean significantly different from zero at the 1-. 5-. and 10-percent levels, respectively.

Table 8: Lee and Strazicich One-Break Unit-Root Tests

	Nominal Interest Rate Differentials			Inflation Rate Differentials			Real Interest Rate Differentials		
Country	T_{B1}	S_{t-1}	***k***	T_{B1}	S_{t-1}	***k***	T_{B1}	S_{t-1}	***k***
Austria	Jun-08	-4.511**	11	Jan-08	-3.955	12	Oct-06	-3.602	12
Belgium	Jun-08	-3.940	11	Apr-08	-3.762	12	Apr-08	-3.553	12
Finland	Feb-08	-3.764	6	Sep-05	-3.000	12	Sep-05	-2.985	12
France	Jul-08	-4.121	12	Apr-05	-3.276	12	Mar-05	-3.476	12
Greece	Feb-09	-3.997	12	Jan-08	-3.601	11	Jun-09	-2.744	6
Ireland	Aug-08	-5.637*	11	Jan-09	-3.544	11	Oct-08	-3.931	11
Italy	Jun-09	-4.553**	11	Aug-05	-3.850	12	Sep-04	-3.606	12
Luxembourg	Apr-06	-3.330	9	Oct-05	-3.655	12	Apr-06	-3.389	12
Netherlands	Aug-08	-3.930	11	Jun-04	-3.573	12	Mar-04	-3.569	12
Portugal	Dec-08	-5.910*	4	May-04	-3.907	12	Dec-05	-3.068	12
Spain	May-08	-3.898	11	Oct-09	-4.043	12	Oct-08	-3.715	12

Note: Critical values for the LM one-break unit-root test statistic in Model C depend somewhat on the location of the break. The 1%, 5%, and 10% critical values for different values of T_{B1}/T come from Table 1 in Lee and Strazicich (2004). The symbols *, **, and † mean significantly different from zero at the 1-. 5-. and 10-percent levels, respectively.

Figure 1: Bilateral Nominal Interest Rate Differentials, January 2001-September 2010

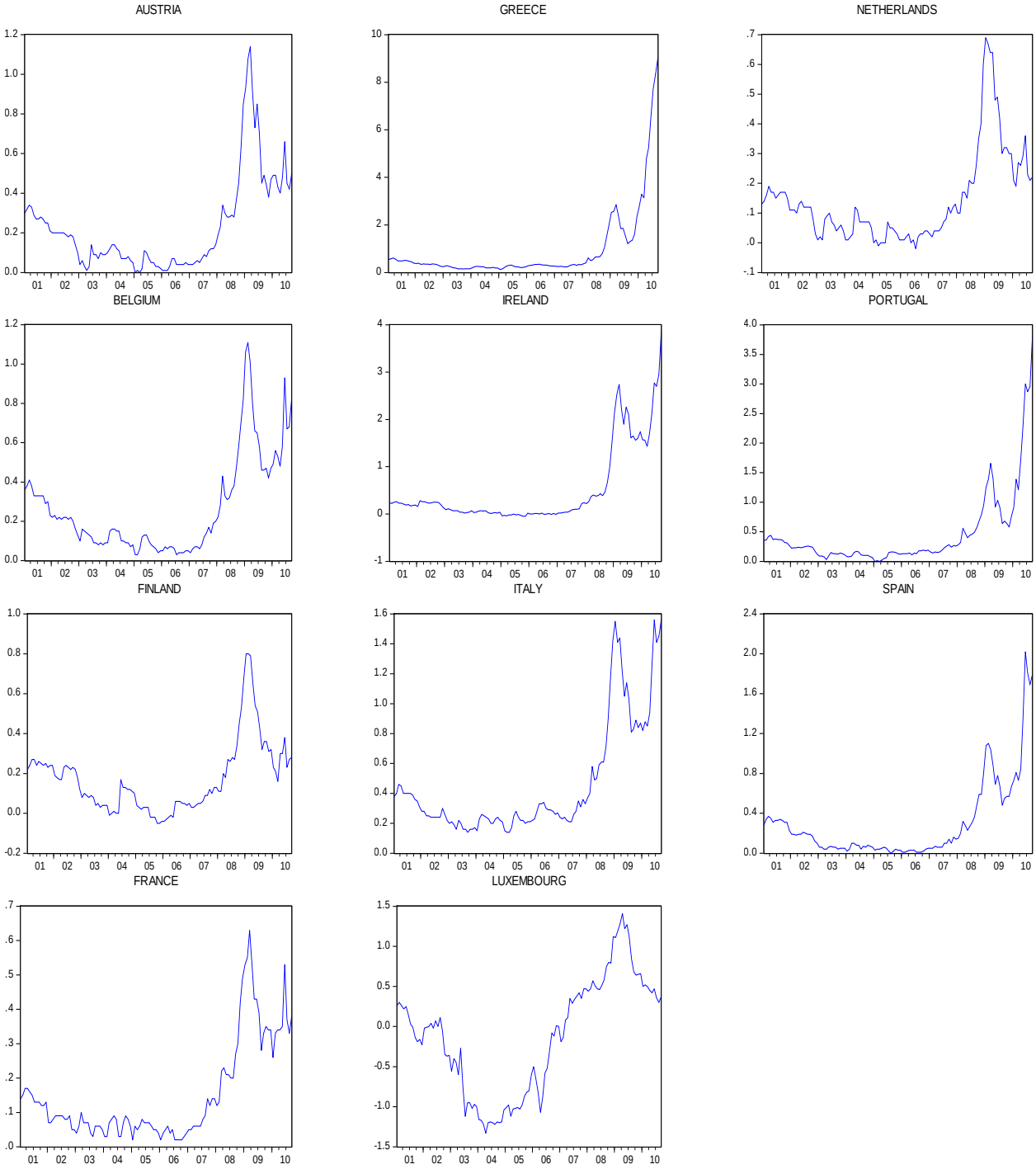


Figure 2: Bilateral Inflation Differentials, January 2001-September 2010

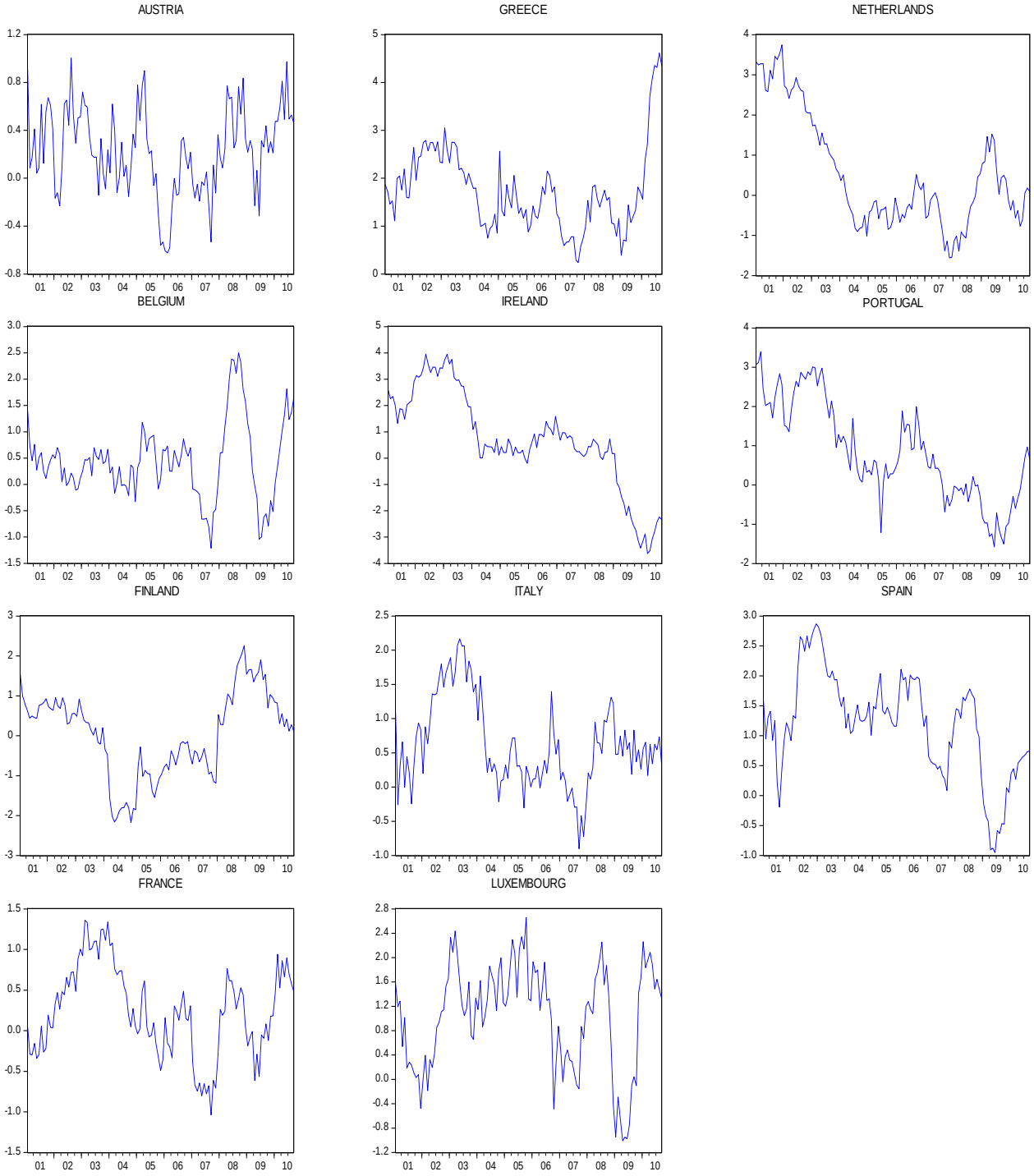


Figure 3: Bilateral Real Interest Rate Differentials, January 2001-September 2010

