STRUCTURAL BREAKS IN LABOR PRODUCTIVITY GROWTH: THE UNITED STATES VS. THE EUROPEAN UNION

2006

Juan F. Jimeno, Esther Moral
and Lorena Saiz

Documentos de Trabajo
N.º 0625

BANCO DE ESPAÑA

STRUCTURAL BREAKS IN LABOR PRODUCTIVITY GROWTH:
THE UNITED STATES VS. THE EUROPEAN UNION

STRUCTURAL BREAKS IN LABOR PRODUCTIVITY GROWTH: THE UNITED STATES VS. THE EUROPEAN UNION (*)

Juan F. Jimeno (**)

BANCO DE ESPAÑA, CEPR AND IZA

Esther Moral and Lorena Saiz (***)

BANCO DE ESPAÑA

(*) We thank seminar participants at Bank of Spain for helpful comments and suggestions. The views in this paper are those of the authors and do not represent the views of the Bank of Spain. Remaining errors are our own responsibility.

(**) juan.jimeno@bde.es.

(***) emoral@bde.es; lsaiz@bde.es.

Documentos de Trabajo. N.º 0625
2006

The Working Paper Series seeks to disseminate original research in economics and finance. All papers have been anonymously refereed. By publishing these papers, the Banco de España aims to contribute to economic analysis and, in particular, to knowledge of the Spanish economy and its international environment.

The Banco de España disseminates its main reports and most of its publications via the INTERNET at the following website: http://www.bde.es.

ISSN: 0213-2710 (print)
ISSN: 1579-8666 (on line)
Depósito legal: M.42034-2006
Imprenta del Banco de España

Abstract

There is a stark contrast between the recent evolution of labor productivity (and TFP) in the US and EU countries. In the US it accelerated around the mid-1990s and there is evidence of reversion to a high-growth regime. In some EU countries, while employment-population ratios started to rise after a period of stagnant employment, labor productivity (and TFP) decelerated.

In this paper we apply univariate and multivariate methods, that have been used to detect structural breaks in productivity growth in the US economy, to EU data to confirm the existence of a significant permanent shift to lower productivity growth in some European countries around the mid-1990s. We find a structural break in mean labour productivity growth in the US around the mid-1990s (towards higher growth), in Continental Europe around the early 1990s (towards lower growth) and no evidence of structural breaks in the UK.

JEL Classification: C22, O47

Keywords: Structural Breaks, Labor productivity, Markov Switching Models.

1 Introduction

Around the middle of the 1990s labor productivity (and Total Factor Productivity, TFP henceforth) accelerated in the United States, and there is some evidence of reversion to a "high-growth regime" (Hansen, 2001, Kahn and Rich, 2004). In stark contrast, around the same time, labor productivity (and TFP) decelerated in some European Union countries. For instance, in the European Union (EU-15) as a whole, the annual rate of growth of real GDP per hour worked declined from an average of 2.3% in the 1980-94 period to 1.4% since the mid-1990s. As shown in Table 1, this deceleration took place in the five largest EU countries and it was particularly acute in Italy and in Spain.

Whether this deceleration in labor productivity growth in the EU is a temporary or a permanent phenomenon is controversial.[1] Some pundits argue that in the EU labor productivity growth has not yet shown an acceleration, as a result of the technical progress brought up by the introduction of new technologies as in the US, because the introduction of new technologies takes time to be translated into higher productivity and investment in these technologies occurred in the EU with some delay with respect to the US. When focusing on the sectorial sources of labour productivity growth, the bulk of the divergence in labor productivity growth between the US and the EU arises mainly from the disappointing performance of non-ICT producing sectors in the latter (Van Ark, et al. 2003), which suggests that the lack of technical progress in ICT is not the main reason for the lagging EU performance in this regard, and that there could be some true in the argument pointing out to delays in the transmission of new technologies to non-ICT producing sectors. There is also the view that the deceleration of labor productivity growth in the EU was a transitory phenomenon due to the fall in capital accumulation during the transition to a balanced growth path with a higher employment rate. A slightly modified version of this view relies on composition effects: increasing employment in Europe brought back into employment low-skilled workers, which had a temporary negative effect on productivity growth. However, although it is true that capital-labor ratios increased by less in Europe to some extent (see Table 1), the decline in labor productivity growth was also due to the fall in TFP growth (with the exceptions of France and the UK).[2]

Nevertheless, as time goes by and there are no signs of recovery, it seems that there could be more fundamental, structural reasons behind the deceleration in productivity growth in the EU. Among them, several hypothesis are put forward:[3] i) a composition of human capital biased towards specific skills

[1] Although there are some relevant measurement issues, they can only explain a minor part of both the acceleration of US labor productivity and the deceleration of European labor productivity.

[2] As for composition effects, they can only explain a small portion of the deceleration of labor productivity growth. Notice, moreover, that, as unemploymet rates are still high in some EU countries, particularly among the low-skilled, the difference in labour productivity *levels* between the US and the EU should widen, were employment rates continuing increasing in the EU.

[3] For an account of these hypotheses, regarding the UK case, see Basu et al. (2003).

which cannot be used to exploit complementarities (Wasmer 2003, Krueger and Kumar, 2004) ii) rigidities that preclude either the introduction of new technologies or the aside investments needed to exploit its full potential, iii) inability to innovate, when growth cannot be sustained any longer through catching up and imitation with respect to the US (Blanchard, 2004).

In this paper we want to confirm the perception that there was indeed a shift to lower productivity growth in some European countries around the mid-1990s. Most studies tend to conclude that in the US economy around that time there was a structural break, towards higher productivity growth. However, as for European countries, there are not many results on the time series properties of productivity in the last decades. Since the detection of structural breaks in economic variables crucially depends on maintained hypotheses about the dynamics of the variable into question, we apply a wide variety of statistical methods designed to detect structural breaks, such as the univariate methods surveyed in Hansen (2001), and multivariate methods, as in Kahn and Rich (2004), that exploit co-movements among macroeconomic variables.

The paper follows in four more sections. Section 2 describes some differences and peculiarities of the evolution of employment, hours worked and, hence, labor productivity in EU countries with respect to the US and present a summary of our results with regard to the dating of structural breaks. Section 3 explains, with some detail, the univariate statistical methods used to detect structural breaks in labor productivity growth in these countries. Section 4 does the same regarding multivariate methods. Finally, Section 5 contains some concluding remarks.

2 Labor input and labor productivity: a preliminary comparison and a summary of results

There are some relevant differences between the US and the five largest EU countries regarding the evolution of labor input and productivity over the last decades. In the US employment growth fluctuated more than in the EU, without any significant trend. By contrast, in some EU countries, notably Spain, Italy, and France, there is a noticeable recovery in employment growth rates after the recession of the early 1990s that resulted in a significant rise in the employment rates of these countries since the mid-1990s. Splitting tentatively the sample in two periods, 1981:1-1994:4 and 1995:1-2004:4, average employment growth was higher in the US in the first sub-period (1.6%) than in the second one (1.2%), while the contrary happened in the European countries [4]. A similar pattern took also place with regard to total hours worked, that shows some deceleration in the US, while in the EU countries they grew by more in the second period

[4] The corresponding figures for the two periods, 1981:1-1994:4 and 1995:1-2004:4, are, respectively: 0.2%, 1.0% in the UK; 0.2%, 0.3% in Germany; 0.1%, 1.0% in France; 0.2%, 0.8% in Italy; and 0.8%, 2.5% in Spain.

than in the first one[5]. This suggests that the growth paths followed by the US and the EU countries are rather dissimilar: while the US can be thought as a country on a balanced growth path, probably hit by a positive technological shock and hence moving to a steady state with higher capital accumulation, the EU countries seem to be in the transition to a growth path characterized by a higher employment rate. This needs to be taken into account when using multivariate methods relying on co-movements among macroeconomic variables to identify shifts in productivity growth, an issue further discussed in Section 4.

Figures 1 to 3 plot growth rates of employment, hours worked and two measures of labor productivity (GDP per worker and GDP per hour worked).[6] As for labor productivity growth, whatever the series used, GDP per worker or GDP per hour worked, growth rates have been significantly higher in the US than in the five-largest EU countries since the mid 1990s. Regarding GDP per worker, average growth rates for the periods 1981:1-1994:4 and 1995:1-2004:4 were, respectively, 1.6% and 2.6% in the US; 1.9% and 1% in Germany; 2.1% and 1.8% in the UK; 1.8% and 1.1% in France; 1.7% and 0.8% in Italy; and 1.9% and 0.8% in Spain. A roughly similar picture arises by looking at GDP per hour worked. The average growth rates during 1981:4-1994:4 and 1995:1-2004:4 were, respectively, 1.7% and 2.8% in the US; 2.8% and 1.6% in Germany; 2.3% and 2.1% in the UK; 2.5% and 2.2% in France; 2.1% and 1% in Italy; and 2.7% and 0.8% in Spain. These simple comparisons suggest the existence of a positive permanent shift in productivity growth in the US, a negative one in Germany, Italy, Spain, and, maybe, in France, and, plausibly, no shift in the UK.

In the rest of the paper we use several statistical methods available to detect structural breaks in economic series and present our results regarding the estimation of break dates in labor productivity growth. The analysis of the existence of structural breaks in the dynamics of economic variables needs to start with the specification of, first, the time series process for the variable into question, and, second, the characteristics of the break to be detected. And the results typically hinges upon these maintained assumptions. For this reason, we specify a wide variety of dynamic processes for labor productivity and estimate break dates and perform statistical tests for structural breaks in each of these specifications.

As a preview of the results, Table 2 presents a summary of the estimated break dates, referred to breaks in the mean of the labor productivity growth. For the US, we typically find two breaks: one towards a lower productivity growth regime in the early 1970s and another to a higher productivity growth regime in the mid 1990s. For Continental European countries, we typically find a permanent decline in mean labor productivity growth around the early 1990s. Finally, for the UK there is no statistically significant evidence of breaks in mean labor productivity growth.

[5] In this case, the US average growth rates in 1981:1-1994:1 and 1995:1-2004:4 were, respectively, 1.6% and 1.0% ; in the UK, 0.0% and 0.7%; in Germany, −0.7% and −0.3%; in France, −0.6% and 0.0%; in Italy, −0.2% and 0.7%; in Spain, 0.2% and 2.4%.

[6] Sources of data and details about the construction of the series are given in the Appendix.

3 Searching for structural breaks (I): Univariate analysis

We start by using univariate statistical methods to identify permanent shifts in mean labor productivity growth in the five largest EU countries and the US. We first assume that productivity growth follows a simple auto-regressive process and perform testing of structural breaks and estimation of break dates. In a second step we also consider alternative specifications, such as long-memory processes with structural breaks and Markov-Switching Regime Models.

3.1 Testing

We first assume for simplicity that labor productivity growth, Δy_t, follows an AR(1) process:

$$\Delta y_t = \alpha + \rho \Delta y_{t-1} + u_t \tag{1}$$

being u_t a white noise process of variance σ^2. Thus, the dynamic properties of labor productivity growth would vary whenever any of the three parameters, α, ρ, σ^2, changes. We focus on structural breaks in long-run labor productivity growth, that is, permanent shifts in the constant or the auto-regressive parameter (α and ρ).[7] As measure of labor productivity growth, we take the quarterly growth rate of GDP per hour worked. First of all, we test the null hypothesis of inexistence of structural breaks (stability in the regression parameters). We consider both individual and joint tests for instability of the intercept and of the auto-regressive parameters. In this context, it is important to remark that we have considered the most simple model (i.e. AR(1)) in order to perform structural breaks tests. Assuming and estimating more complex models could undermine the power of the tests[8].

We have performed three tests[9]: i) Nyblom's L test (Nyblom, 1989, and Hansen, 1992) that has locally optimal power and does not require knowing a priori the date of the break, ii) the Quandt Likelihood Ratio (QLR) (Quandt, 1960) or maximum Wald statistic (Sup-W) and iii) the logarithm of Andrews-Ploberger exponential Wald statistic (Exp-W). The Sup-W and Exp-W tests check for structural breaks making the assumption of unknown break date.[10]

[7] In this context, Hansen (2001) surveys methods available to perform i) Tests for a structural breaks of unknown timing, ii) Estimation of the timing of a structural break, and iii) Tests to distinguish between a random walk and broken time trends.

[8] Nevertheless, additionally, we have performed all these tests with the model selected by the Bayesian Information Criteron and the results do not change significantly. We do not include them in the paper but they are available upon request.

[9] Other tests of parameter instability are the famous Chow test (Chow, 1960) whose main disadvantage is that the break point must be a priori known, and the CUSUM tests of recursive residuals and of squares of recursive residuals of Brown, Durbin and Evans (1975) which have poor asymptotic power. We have also performed CUSUM and CUSUM SQ tests and in all cases we find that they reject parameter stability.

[10] The former test is the largest value of all the sequence of Wald F-statistic calculated for each date and the latter is the exponential transformation of that sequence of F-statistic.

Results, collected in Table 3, suggest that there is statistical evidence of instability in both parameters, and, consequently, in mean labor productivity growth, in Spain, Italy, Germany and France. In the case of United States and United Kingdom only the Nyblom's test rejects the null of joint stability of both the intercept and the auto-regressive parameter.[11]

In a second step, we search for multiple break dates using several procedures developed by Bai and Perron (1996, 2003a, 2003b): i) a test of the null of no break against the alternative of the existence of a fixed number of breaks (one, two and three), ii) a test of the null of no break against an unknown number of breaks (double maximum tests), and finally, iii) a test of the null of k breaks against the alternative of $k+1$ breaks (for $k = 1, 2$). The two panels of Table 4 report the results, the top panel regarding a pure structural change AR(1) model, that is, considering changes in the intercept and in the AR parameter, and the bottom panel regarding a partial structural change AR(1) model in which only changes in the intercept are considered.

Results suggest that there is at least one structural break in any of the parameters in Germany, Italy and Spain. For these three countries, the first two tests reject the null of no break in both the general and the partial models, while the third test accepts the existence of two breaks in Germany and Italy, and one in Spain. In the case of France we only find evidence in favor of one breakpoint in the partial structural change model. In general, there is no evidence of structural breaks for the United Kingdom. And, finally, the three tests for both models indicate that labor productivity in the United States would have experimented at least one break.

3.2 Dating

Under the assumption of linearity and homoskedasticity of the covariance matrix, the natural candidate for the estimation of the break date is that corresponding to the largest value of Wald test sequence (sup-W), that is algebraically identical to the date that minimizes the sum of squared residuals. In general, for practical purposes, it is preferable and more efficient estimating regression models by OLS splitting the sample at each possible break date, and, then, finding the one that minimizes the full-sample sum of squared errors sequence.[12]

In Table 5 we report the estimated breakpoints and values for the mean productivity growth. We include the point estimation of the break date and the confidence interval at 90% for the five largest EU countries and United

Both statistics have unknown distributions. Andrews (1993) and Andrews and Ploberger (1994) provided the corresponding critical values, and Hansen (1997) developed a method to calculate p-values.

[11] We also tested for variance stability. For the US and the UK we found that there seems to be shifts in the variance of the AR(1) process for labor productivity around 1983 and 1985, respectively. It should be noticed that under variance instability, the tests we performed for stability of the intercept and the auto-regressive parameter have low power and they would yield incorrect results (see Cogley and Sargent, 2005).

[12] Bai (1994,1997a and 1997b) derives the asymptotic distribution of the breakdate estimator and shows how to construct confidence intervals.

States. For Italy and Germany, as the previous tests revealed the presence of two breaks in the mean productivity growth, we have estimated two instead of one. In both cases the first break date would be around the end of the 1970s, towards lower mean growth, and the second one at the beginning of the 1990s, to an even lower mean growth rate. The shift to a lower mean growth regime in Spain would have happened in the mid-1980s, while in France it would have taken place at the onset of the 1990s[13]. For United States, we have performed estimation of two break dates finding that the first one would be around 1973 and the second one between 1996 and 1997. This confirms the results obtained in previous studies: Hansen (2001) finds strong evidence of these two breaks for labor productivity in the US manufacturing/durables sector, Kahn and Rich (2004) obtains similar breaks in US non-farm business sector output per hour growth, and Benati (2005) achieves the same results for three series of output per hour growth for business, non-farm business and manufacturing sectors. In the case of the United Kingdom, we have estimated a break date even though we did not find evidence of a structural change in mean productivity growth. Not surprisingly, we obtain a very uncertain outcome, with rather large confidence intervals around the estimated break date.[14]

An alternative approach at estimating break dates is to build upon the results of the Structural Break Augmented Fractional Dickey-Fuller (SB-AFDF) test[15], that we explain in more detail in section 3.3. The estimated break dates are presented in Table 6. Under this approach, the break is assumed to have occurred at an unknown date, but by performing the test in a sequential way, the break date can be identified as that corresponding to the minimum statistic. This approach yields results consistent with those of previous tests for the United States, France and Spain. As for Germany and Italy, it signals to the existence of two break dates, one at the end of 1970s and another during the 1990s. Lastly, for United Kingdom the break would have taken place at the beginning of the 1980s, but in this case it is not statistically significative.

Summing up, for France, Germany and Italy we detect an important slowdown in the mean labor productivity growth around the mid-1990s. For Spain this slowdown would have occurred before, in the mid-1980s. United States productivity growth would have experimented a decline around 1973, and then a recovery to higher mean productivity growth around 1997, reaching an average rate even greater than the one observed before 1973. And finally, we do not find

[13] This result coincides with the findings of Beffy and Fourcade (2005).

[14] This also has to do with the instability in variance shown by the labor productivity growth in the UK. In the presence of variance shifts, the tests we have used would become invalid. However, all the time in our analysis we have controlled by heteroskedasticity and serial autocorrelation using the Newey-West variance-covariance matrix.

[15] Another way of testing for the presence of structural breaks is using unit root tests (see Perron, 1989, Andrews and Zivot, 1992). Perron (1989) provides a test of a difference-stationary process with breaks versus the alternative of a trend-stationary process with breaks, treating the break dates as known a priori. Andrews and Zivot (1992) propose a modification of ADF unit root tests with endogenous break date, but only under the alternative of a trend-stationary process. We have performed Andrews and Zivot's test for the labor productivity series and in the majority of cases the test is rejected. The results are not included for brevity reasons but they are available upon request.

any clear evidence in favour of a structural break in labor productivity growth in the case of the UK.

3.3 Long-memory processes and structural breaks

An alternative approach for explaining dynamic persistence in economic variables is to consider fractional integration. There is a connection between fractional integration and common notions of structural change. Stochastic processes with short memory or stationarity that exhibit structural changes can display similar characteristics (such as auto-correlogram function, periodogram, etc) as the ones observed in long memory or long range dependent processes. Dolado, Gonzalo and Mayoral (2002, 2005) develop a time-domain test in order to distinguish between long-memory and structural breaks. In essence, it is an ADF test, but in this case the null is that the process is integrated of order d ($I(d)$), with $0 < d < 1$. The alternative hypothesis entails that the process is $I(0)$, but allowing for structural breaks (in the intercept and in the deterministic trend in case of incorporating deterministic components) at unknown dates. This distinction is relevant because the current shocks will have temporary effects of greater (integrated process) or lower duration (stationary process), and only occasional events or shocks associated to structural breaks will have permanent effects on the long-run level of the series. So, the distribution of the durations of the shocks determines whether or not the process is fractionally integrated. Fractional integration requires that a small percentage of the shocks have long durations to generate high-order auto-correlations. And when we observe stationary series with structural breaks, these few occasional shocks with long-duration effects would produce certain persistence or symptoms of non-stationarity in the series.

Results from applying this test to labor productivity are summarized in Table 6.[16] We have considered several cases of fractional integration, with values of d, the order of integration, lower and higher than 0.5, as the SB-AFDF statistic follows different distributions depending on the value of this parameter: for values of d smaller than 0.5 the series will tend to be stationary (this means the expected number of surviving shocks after several periods will be small) and the SB-AFDF statistic will be normally distributed; alternatively, when d takes values greater than 0.5, the process will be more non-stationary (higher number of expected surviving shocks) and the SB-AFDF follows a special distribution. We have considered two models: i) a pure structural change model or the crash and changing-growth model in Perron's terminology in which changes in the deterministic trend as well as in the intercept are allowed, and ii) a partial structural change model with shifts in the constant or the crash model. We have performed the estimation of the SB-AFDF statistic in a sequential way for each point without imposing a break date a priori. Then, we have chosen the infimum of that sequence to develop the test.

According to the results, in all countries but the United Kingdom it is possible to accept the alternative of stationarity with structural breaks for several

[16] We are very grateful to Juan J. Dolado, Jesús Gonzalo and Laura Mayoral for their help in performing the test and for providing us the programs and routines in MATLAB.

values of d and for both models. However, for the British case, we are able to reject the null of fractional integration only for $d = 0.9$ under the first model, suggesting that productivity growth follows a long-memory process[17]. Besides, notice that it is the only country for which we are not able to reject the hypothesis of fractional integration with high d values (greater than 0.5). This result indicates that UK labor productivity growth, apart from having a high persistence, would exhibit a marked instability in variance (i.e. process non-stationary in covariance).

3.4 Markov-switching productivity regimes

A flexible extension of the AR(1) process is Hamilton's (1989) model that allows for gradual changes and transition periods between different regimes. The simplest specification of this model is the following Markov process with two regimes distinguished by different mean growth rates:

$$\begin{gathered} \Delta y_t = \mu_{S_t} + \phi(\Delta y_{t-1} - \mu_{S_{t-1}}) + \varepsilon_t \\ Var(\varepsilon_t) = \sigma^2 \\ p = \Pr[S_t = 1/S_{t-1} = 1] \\ q = \Pr[S_t = 0/S_{t-1} = 0] \end{gathered} \tag{2}$$

where Δy_t is labor productivity growth, ϕ is an auto-regressive coefficient, μ is mean productivity growth and ε_t is a Gaussian white noise with variance σ^2. S_t is the state variable, that takes values 0 or 1 depending on the two states of nature, low or high mean growth, with transition probabilities of remaining in the same state of q and p, respectively. The parameters for each state and the state at which the economy is at every moment can be simultaneously estimated, without any need to assume the break date nor the nature of the change (gradual or abrupt)

We have estimated a MS-AR(1) model with changing mean for Spain and Germany. For Italy and France we estimate a MS-AR(4) model, that supplies the most reasonable results in terms of goodness of fit, parameter values and probabilities. Here, in contrast with the before sections in which we were mainly interested in testing, the model specification is crucial to obtain reliable results. Table 7 summarizes the estimated models and Figure 4 plots the smoothed state probabilities. We find that in several European countries the mean duration of the low-mean growth regime is similar than that of the other regime, and the ergodic probability of remaining in that state is around 0.5. This would indicate us that in all these four countries there is evidence of a switch to a low productivity regime more or less in the middle of the sample, this is, sometime along the 1980s, earlier in Spain, Germany, and Italy, and later in France. The most intense productivity slowdown would have happened in Spain and Italy, and the smoothest would be that of France. As for the UK and the US,

[17] We have also estimated the order of fractional integration d for the british series by means of the generalized minimum distance estimator proposed by Mayoral (2006), and we obtained a value of 0.83.

the estimation of this specification for labor productivity growth yields non-satisfactory results. In fact, for these countries, a MS-AR(1) with only variance switching, namely,

$$\begin{array}{c} \Delta y_t = \mu + \phi(\Delta y_{t-1} - \mu) + \varepsilon_t \\ Var(\varepsilon_t) = \sigma^2_{S_t} \\ p = \Pr[S_t = 1/S_{t-1} = 1] \\ q = \Pr[S_t = 0/S_{t-1} = 0] \end{array} \tag{3}$$

seems to work better. Results are also reported in Table 7 and Figure 4. There seems to be a decline in the variance of labor productivity growth around 1989 for the UK and around 1983 for the US.[18] The results from estimation of a MS process combining both mean and variance switching for these two countries were also non-satisfactory.

3.5 Summary of results

Productivity growth series display quite a noisy behavior at quarterly frequencies. This together with data limitations generate not very consistent results from alternative univariate structural breaks tests. In any case, as a summary of results within the univariate framework, we draw the following list of findings:

i) There is some statistical evidence of a slowdown in mean labor productivity growth for several European countries, but the United Kingdom. However, the number, size, and dating of breaks are not really clear, as results depend on the statistical models and tests considered. In fact, the estimated confidence intervals for breaking dates are too wide in many cases. Moreover, for Italy and Germany we obtain conflicting results regarding the number of breaks: in some case we obtain evidence of two breaks, one in the seventies and another in the nineties, but by estimating a Markov Switching model, we are led to conclude that there was only one break towards lower mean growth at the mid-seventies.

ii) As for the US, in line with other works in the literature, we detect a productivity growth slowdown in 1973 and an important resurgence around the mid-nineties. We also find evidence of variance instability.

iii) Regarding the UK, we also find evidence of variance instability and symptoms of fractional integration instead of structural breaks.

4 Searching for structural breaks (II): Multivariate analysis

Separating permanent and transitory shifts in labor productivity growth is a hard task, especially when the shifts take place at the end of the sample period and, hence, subsequent data to confirm the nature of the shifts are not yet available. Hence, not very surprisingly, univariate statistical methods applied

[18] This last result confirms previous findings regarding US GDP growth (see McConnell and Perez-Quiros, 2001)

to detecting structural breaks in productivity growth do not always yield robust results, especially in the case of very volatile series at high frequencies.

As an alternative approach, we rely on co-movements among several related macroeconomic variables to identify shifts in the long-trend of the variable of interest. Within this framework, there are three alternative approaches: i) using a large number of variables to estimate factor models, ii) exploiting restrictions implied by economic theory on these co-movements, which allows to specify low dimensional systems, and iii) using cross-country restrictions (i.e. simultaneity) on the dating of structural breaks.

Here we follow the second and third approaches. In the case of labor productivity growth, the neoclassical growth model establishes that, under a balanced-growth path, both labor and capital income shares, on the one hand, and employment rates and work effort per capita, on the other, are constant. This implies that labor productivity, consumption per capita, wages and the capital-labour ratio have a common stochastic trend, which can be associated to long-trend productivity growth. Below we exploit these restrictions in the estimation of structural break dates.

4.1 Estimating breaks under common stochastic trends

If indeed labor productivity, consumption per capita, wages, and the capital-labour ratio have a common stochastic trend, they will be co-integrated. After normalizing the series by hours, preliminary co-integration analysis shows that only France and Germany exhibited at most two co-integration relations among the variables considered, as it is the case in the US. In Italy, we only find one co-integration relation, and there is no evidence of co-integration in Spain and the UK . But it is important to signal that conventional co-integration tests could yield misleading results in the presence of structural breaks (Leybourne and Newbold, 2003). The study of co-integration relations in presence of structural breaks is very complicated and even more, in case of more than two series, since the instability or break could be in all or several co-integration coefficients, in the error correction coefficient, in all long-run relations or in some of them, or in the short-run dynamics. And the breaks could happen contemporaneously in all the series or not. Several tests have been recently developed in the statistical literature for testing co-integration in presence of breaks (e.g. residual-based co-integration test with regime shifts of Gregory and Hansen, 1996; co-integration rank tests of a VAR with level shift at unknown time, as in Lütkepohl, Saikkonen and Trenkler, 2004; and panel co-integration tests with multiple structural breaks,as in Westerlund, 2006; among others) but there are no consensus about which is more powerful and offers more robust results. Instead, we have carried out recursive eigenvalues test (Hansen and Johansen, 1999) to detect instability in the number of co-integration relations or co-integrating rank (probably, due to the presence of structural breaks in the co-integration relation), and we found symptoms of a slight instability in the co-integration relations for Spain, Italy, France and United Kingdom. Thus, the results commented below, which are obtained under the maintained hypothesis of co-integration, should be taken with some caution for these countries.

To detect common structural breaks we apply the methodology developed by Bai et al. (1998) to a system consisting of labor productivity, consumption per hour and labour compensation per hour. Bai et al showed that there seems to be important gains in terms of precision by using multivariate inference about break dates since the width of the confidence interval is inversely related to the number of series which have a common break date. And even if series with no common breaks were included, the width of the interval for the break date does not increase. This test is similar to those of the univariate case since it consists in calculating sequences of F-Wald tests, but here we check for the null of a common break date in the intercept treating the break date as unknown.

We estimate unrestricted bivariate and trivariate VAR models for the series in first differences and one lag[19]. Table 8 reports the results of these tests, as well as the estimated common structural break dates (and their confidence intervals) for the bivariate VAR (panel A) and trivariate VAR (panel B). It turns out that it is possible to find a statistically significant structural break in labor productivity for all European countries, except for the UK, although at different dates. In Italy a structural break common to the three series considered is found exactly in the middle of the 1990s, while in the rest of the countries the break date is found to be previous. In particular, the common break date is estimated to have taken place around mid-1980s in Spain, at the end of that decade in France and finally, in the second half of the 1970s and at the beginning of the 1990s in Germany. For the US, we detect two break dates, one in the early 1970s and another one around the mid-1990s. Overall, the confidence intervals are reduced when compared to the estimation of break dates with univariate methods.

The next step consists in estimating a VECM model but allowing for a break in the growth rate of the common stochastic trend (i.e. shift in the intercept term), as in King, Plosser, Stock and Watson (1992) and Bai et al. (1998):

$$\Delta Y_t = \mu + \lambda d_t(k_0) + A(L)\Delta Y_{t-1} + \beta\alpha' Y_{t-1} + e_t \tag{4}$$

where ΔY_t is the vector of variables (i.e. labor productivity, consumption per hour and compensation of employees per hour) in logarithms and first differences, α is the co-integration vector, $d_t(k_0)$ is a dummy variable which takes value 0 for $t < k_0$ and 1 otherwise, and k_0 is the a-priori unknown breakpoint. In spite of the weak results of co-integration obtained, we think it is interesting to incorporate the information of the possible long-run relations with the aim of gaining some precision in the statistical inference and reducing the width of the confidence intervals.

We estimate the co-integration vector by dynamic ordinary least squares (DOLS) and then one VECM for each possible break date. Results are summarized in Table 9. In general, they are similar to those obtained from the

[19] We have also performed all these tests with the model selected by the Bayesian Information Criteron and the results do not change significantly. We do not include them in the paper but they are available upon request.

unrestricted VARs, except for two cases: the United Kingdom, where we obtain statistical evidence of a break in the second half of the 1980s, and Spain, where now the break is found at the beginning of the 1990s. For this last one, it is interesting that when we re-estimate the unrestricted VAR starting the sample from 1986 onwards (i.e. after the severe industrial restructuring process suffered by Spain), we obtain the same break than in the VECM (see summary in Table 2). Perhaps the great magnitude of the structural change experimented by the Spanish economy in the mid-eighties could have hidden other relevant breaks such as that of 1993.

With all, the most important thing is that in all cases the width of the confidence interval in the VECM is narrower than in the VARs, showing the gains of including the co-integration vectors in the system. However, we again insist that all of these results must be taken carefully, given the instability of co-integration relations for European countries. For the United States, the co-integration vectors are very stable and when we test for unit coefficients, that is, (1,-1,0) and (1,0,-1), the null is not rejected at 5% significance level. Thus, we include the results of the VECM imposing this restriction. They are in line with the results of VECM with estimated co-integration vectors, but, not surprisingly, the width of the confidence interval is smaller.

4.2 Exploiting cross-country restrictions

A possibility to be exploited in the estimation of break dates is that similar countries may have experienced them simultaneously.

When performing the test of common break dates for the labor productivity growth series of all the Continental European countries considered (see Table 10), the null hypothesis is not rejected at the 5% significance levels and the break date is estimated around the second half of the 1990s. When Spain is not included (see footnotes in Table 10), there is again statistical evidence supporting the common break date in the 1990s. And finally, considering only the United States and the United Kingdom, the results do not show any favorable sign of a common shift at the same time. As a complementary exercise we have developed these tests for sets of pairs of similar European countries: Italy and Spain, and France and Germany. In both cases there is not enough statistical evidence in favor of a common break at the mid-1990s.

4.3 Markov switching productivity regimes: Multivariate analysis

As in the univariate case, Markov processes could be used to analyze changes in productivity growth. For the US this has been done by Kahn and Rich (2004), who estimate the following multivariate model:

$$\begin{aligned}
\Delta y_{i,t} &= \gamma \Delta \Pi_t + \lambda_i \Delta x_t + \Delta z_{it} \\
\Delta \Pi_t &= \mu_0(1 - S_{1t}) + \mu_1 S_{1t} + \phi_1 \Delta \Pi_{t-1} + v_t \\
x_t &= \tau S_{2t} + \phi_{11}^* x_{t-1} + \phi_{12}^* x_{t-2} + \varepsilon_t \\
z_{i,t} &= \psi_{i1} z_{i,t-1} + \eta_{i,t} \\
Var(v_t) &= Var(\varepsilon_t) = 1; Var(\eta_{i,t}) = \sigma_i^2 \\
i &= 1, 2, 3, 4
\end{aligned} \tag{5}$$

where $y_{i,t}$ is a vector composed of the logs of labor productivity, real compensation per hour, private consumption per hour, and detrended hours of work[20], Π_t is the permanent component that is common to all the series, x_t is the transitory component that is also common to all the series (plausibly related to the business cycle), and $z_{i,t}$ is the idiosyncratic component of each series (measurement errors and noise). The parameter γ is the factor loading of the common permanent component, and λ_i are the factor loadings of the common transitory component in each of the series. Kahn and Rich (2004) impose that the factor loadings of the permanent components are the same for the first three series ($\gamma_1 = \gamma_2 = \gamma_3 = \gamma$), and zero for the fourth one ($\gamma_4 = 0$).[21] In order to minimize the number of parameters to estimate, they assume that the common permanent component is differenced stationary, following an AR(1) with two means depending on the state (high or low mean) as proposed by Hamilton (1989), that the common transitory component is an AR(2) with two means (zero and negative) following Friedman's plucking model (Friedman 1964,1993) about asymmetries in business cycles, and finally, that the idiosyncratic components are AR(1) except for the fourth variable that follows an AR(2). Notice that the permanent and transitory regimes are independent and the unit variance of the error terms of both components is an identifying restriction. This specification uses more information than just the one provided by productivity data. Moreover, break dates are not fixed a priori, and its estimation yields dating of regime switches together with the probability of a switch and, hence, an estimated duration of each regime episode. From their results, Kahn and Rich (2004) conclude that labor productivity in the US returned to the high growth regime around 1997.

It seems tempting to replicate this same methodology to the analysis of labor productivity growth in EU countries. We try several specifications (including or not the plucking effect, not imposing restrictions on γ, using only two or three variables, etc) and the results were disappointing. In any case, when performing the corresponding likelihood-ratio tests, we find that transitory regimes are not significant. In general, introducing the restriction in γ does not change the results and the restrictions are not rejected. By contrast, the results change a lot depending on the set of variables considered in the system.

[20] Notice that we have also included in the system detrended hours of work as in Kanh's paper, in order to capture better the business cycle or transitory component. We have employed the popular Hodrick-Prescott filter with $\lambda = 1,600$ to extract the cyclical component of the series of worked hours.

[21] They can impose these restrictions because the series are cointegrated (two cointegration relations) and don't reject unit coefficients (1,-1,0) and (1,0,-1).

Table 11 reports some results. For the US the size of the regime-shift for the high and low growth states, $\gamma(\mu_1 - \mu_0)$, is estimated to be 0.39, which corresponds to an annualized gap of rates of growth between both regimes of 1.55%. For Spain this gap is 0.62 (2.49% in annual terms), for France is 0.32, (1.28%), and for Germany 1.23 or (4.91%). In the cases of United Kingdom and Italy, we do not obtain reliable results from the estimation of this model. With respect the duration of the regimes, for US we find that the low-growth regime would have an expected average duration of 15.6 years with an unconditional probability of being in that regime of 0.30. However, in Spain this regime would last on average 20.8 years while in Germany it would last 16.7 years and in France, 20.8 years, with ergodic probabilities of around 0.50 in the three cases.

More interestingly, Figure 5 plots the smoothed and real-time probabilities of remaining in low-growth state at every date. Among the six countries, in four of them (Germany, Italy, Spain and France) there seems to be a switch to a low-growth regime somewhere between the early and mid 1990s, in the UK there is once again no evidence of a switch, while in the US, labor productivity stayed in the low-growth regime from the early 1970s up to the mid 1990s, when it entered in a new episode of high productivity mean growth.

For France and Italy we observe large differences between smoothed and real-time probabilities (that is, considering or not all the sample to estimate them). As expected, real-time probabilities detect regime-shifts several quarters after smoothed probabilities do, because they do not use the entire sample. This fact would indicate that the regime-changes are so gradual than it is necessary a lot of information to detect them properly.

4.4 Summary of results

Exploiting the restrictions imposed by the economic theory on the co-movements among macroeconomic series and include in our analysis two additional variables related to labor productivity, i.e. compensation of employees and consumption, we obtain important gains in the precision of the statistical inference as the width of the confidence intervals decreases and we are able to detect breaks more clearly. Under this approach, the main findings are:

i) There is supporting evidence of a switch towards a low-mean growth regime between the early and mid-nineties for four European countries (Germany, Italy, France and even Spain) from estimation of unrestricted VARs, VECMs and Dynamic Factor models with Markov Switching . When we exploit the cross-country information, again there is statistical evidence of a simultaneous break for the majority of countries considered around the second half of the nineties.

ii) As for United States, we confirm the results obtained under the univariate approach: a slowdown in productivity growth in the early 70s and the recovery of the mid-1990s.

iii) In contrast, labour productivity growth in the United Kingdom does not show symptoms of any break and the existence of a common break with United States around the mid-nineties is clearly rejected.

5 Concluding Remarks

The fall in labor productivity growth observed in some EU countries is a contentious issue. There seems to be no consensus about neither its cyclical nature nor the transitory and permanent factors that could account for the deceleration in labor productivity. For some, it is just a mere blip associated to the transition to a balanced growth path with a higher level of employment rates. For others, it is an indication of the crude reality arising from the loss of dynamism and innovation in these countries that may have taken place sometime along the 1980s and became more evident when the new technologies started to show all of their potentials at the US.

In this paper, we have not solved this dispute, but we have shown, through a battery of statistical tests, both in univariate and multivariate contexts, that, in fact, something "structural" seems to have happened to labor productivity growth in Germany, Italy, France and Spain, causing a lower long-trend growth since the 1990s. This, together also with the UK experience, labeled as the "missing productivity growth" puzzle by Basu et al. (2003), should foster more research at identifying the causes why the introduction of new technologies has led to different productivity developments at the two sides of the Atlantic.

6 Appendix: Data sources

US series are taken from the Bureau of Labour Statistics. Series for European countries have been obtained using several data sources. In particular, output and employment data come from each country's National Statistical Institute (NSI), and consumption, compensation of employees and GDP deflator data come from Eurostat (except in the case of Germany, for which they were obtained from the Bundesbank). Output for each country refers to its Gross Domestic Product in national currency (in euros in the case of EMU countries for the whole sample period) and in real terms, coming from Quarterly National Accounts, as well as real private consumption. Data on worked hours are taken from the OECD Productivity Database. Population data come from several sources: OECD (for UK and Italy, population aged 15 to 65), the corresponding NSI (for Germany and France, total population) and the Ministry of Economy (Spain, population over 16 years of age).

All series had originally a quarterly format, except hours data, which are defined as annual hours actually worked per person in employment. Thus, we applied a linear interpolation in order to get a quarterly series for hours worked.

The length of the sample period is different for each country and only in the case we make a multi-country analysis, we use the sample period which is common to all countries: from 1980Q1 to 2004Q4. In the rest of the paper, we make use of the broadest time period for which data for all series are available in each country, that is, from 1947Q1 in the case of USA, from 1960Q4 for Germany, from 1970Q1 for Italy, from 1971Q2 for UK, from 1978Q1 for France and finally, from 1980Q1 for Spain.

References

[1] ANDREWS, D. W. K. (1993). "Tests for Parameter Instability and Structural Change with Unknown Change Point", *Econometrica,* Vol. 61, No. 4, pp. 821-856.

[2] ANDREWS, D. W. K., and W. PLOBERGER (1994). "Optimal Tests when a Nuisance Parameter is present only under the Alternative", *Econometrica,* Vol. 62, No. 6, pp. 1383-1414.

[3] ANDREWS, D. W. K., and E. ZIVOT (1992). "Further Evidence of the Great Crash, the Oil-Price Shock and the Unit-Root Hypothesis", *Journal of Business and Economic Statistics,* Vol. 20 (1), pp. 25-44.

[4] BAI, J. (1994). "Least Squares estimation of a shift in linear process", *Journal of Time Series Analysis.*

[5] — (1997a). "Estimation of a change point in multiple regression models", *The Review of Economics and Statistics,* Vol. 74, pp. 551-563.

[6] — (1997b). "Estimating multiple breaks one at a time", *Econometric Theory,* Vol. 13, pp. 315-352.

[7] BAI, J., L. LUMDSDAINE and J. H. STOCK (1998). "Testing for and dating common breaks in Multivariate Time Series", *The Review of Economic Studies,* Vol. 65, 3, pp. 395-432.

[8] BAI, J., and P. PERRON (1998). "Estimating and Testing Linear Models with Multiple Structural Changes", *Econometrica,* Vol. 66, pp. 47-78.

[9] — (2003a). "Computation and Analysis of Multiple Structural Change Models", *Journal of Applied Econometrics,* Vol. 18, pp. 1-22.

[10] — (2003b). "Critical values for multiple structural change tests", *The Econometrics Journal,* 6, June.

[11] BASU, S., J. FERNALD, N. OULTON and S. SRINIVASAN (2003). *The case of missing productivity growth: Or, does the information technology explain why productivity accelerated in the US but not in the UK?,* NBER, Working Papers No. 10010.

[12] BEFFY, P., and N. FOURCADE (2005). "The slowdown in labor productivity in the 1990s: the impact of the employment policies", *Économie et Statistique,* No. 376-377, June.

[13] BENATI, L. (2005). "Drift and breaks in Labour productivity", paper presented at CEPR/EABCN Conference: *Productivity and the Business Cycle: Evidence for Europe and the U.S,* mimeo, Helsinki, 28-29 November.

[14] BLANCHARD, O. (2004). "The economic future of Europe", *Journal of Economic Perspectives,* Vol. 18, No. 4, Fall, pp. 3-26.

[15] BROWN, R. L., J. DURBIN and J. M. EVANS (1975). "Techniques for Testing the Constancy of Regression Relationships Over Time", *Journal of the Royal Statistical Society,* Series B, Vol. 37, pp. 149-192.

[16] CHOW, G. C. (1960). "Tests of Equality between Sets of Coefficients in Two Linear Regressions", *Econometrica,* Vol. 28, 3, pp. 591-605.

[17] COGLEY and SARGENT (2005). "Drift and volatilities: monetary policies and outcomes in the post WWII US", *Review of Economic Dynamics,* Vol. 8, pp. 262-302.

[18] DOLADO, J. J., L. MAYORAL and J. GONZALO (2002). "A Fractional Dickey Fuller Test for Unit Roots", *Econometrica,* Vol. 70, No. 5, September, pp. 1963-2006.

[19] — (2005). *What is What?: A Simple Time-Domain Test of Long-memory vs. Structural Breaks,* mimeo.

[20] FRIEDMAN, M. (1964). "Monetary Studies of the National Bureau", *The National bureau enters its 45th year, 44th Annual Report,* pp. 7-25. Reprinted in Milton Friedman (1969), *The Optimum Quantity of Money and Other Essays,* Chap. 12, Chicago: Aldine, pp. 261-284.

[21] — (1993). "The 'plucking model' of Business fluctuations revisited", *Economic Inquiry,* April, pp. 171-177.

[22] GREGORY, A. W., and B. E. HANSEN (1996). "Residual-based tests for Cointegration in Models with Regime and Trend Shifts", *Journal of Econometrics,* Vol. 70 (1), pp. 99-126.

[23] HAMILTON, J. D. (1989). "A New Approach to the Economic Analysis of Nonstationary Time Series and the Business Cycle", *Econometrica,* Vol. 57, No. 2, pp. 357-384.

[24] HANSEN, B. E. (1992). "Testing for Parameters Instability in the Linear Models", *Journal of Policy Modeling,* Vol. 14, 4, pp. 517-533.

[25] — (1997). "Approximate Asymptotic p-values for Structural Change Tests", *Journal of Business and Economic Statistics,* Vol. 15 (1), pp. 60-67.

[26] — (2001). The New Econometrics of Structural Change: Dating Breaks in U.S. Labor Productivity, *Journal of Economic Perspectives,* Vol. 15, 4, pp. 117-128.

[27] HANSEN, H., and S. JOHANSEN (1999). "Some tests for parameter constancy in cointegrated VAR-models", *Econometrics Journal,* Vol. 2, pp. 306-333.

[28] KAHN, J. A., and R. W. RICH (2004). *Tracking the New Economy: Using Growth Theory to Detect Changes in Trend Productivity,* mimeo, Federal Reserve Bank of New York.

[29] KING, R. G., C. I. PLOSSER, J. H. STOCK and M. W. WATSON (1991). "Stochastic Trends and Economic Fluctuations", *American Economic Review,* Vol. 81, pp. 819-840.

[30] KRUEGER, D., and K. B. KUMAR (2004). "US-Europe differences in technology driven growth: quantifying the role of education", J*ournal of Monetary Economics,* 51, pp. 161-190.

[31] LEYBOURNE, S., and NEWBOLD, P. (2003). "Spurious rejections by cointegration tests induced by structural breaks", *Applied Econometrics,* Vol. 35, pp. 1117-1121.

[32] LÜTKEPOHL, H., P. SAIKKONEN and C. TRENKLER (2004). "Testing for Cointegration Rank of a VAR process with a level shift at unknown time", *Econometrica,* Vol. 72, pp. 647-662.

[33] MAYORAL, L. (2006). *Minimum Distance Estimation of Stationary and Non-stationary ARFIMA Processes,* Working Paper No. 959, Universidad Pompeu Frabra.

[34] MCCONNELL, M., and G. PÉREZ-QUIRÓS (2000). "Output fluctuations in the United States: What has changed since the early 1980's?", *American Economic Review,* Vol. 90, No. 5, pp. 1464-1476.

[35] NYBLOM, J. (1989). "Testing for constancy of parameters over time", *Journal of the American Statistical Association,* Vol. 84 (405), pp. 223-230.

[36] PERRON, P. (1989). "The Great Crash, the Oil Price Shock and the Unit Root Hypothesis", *Econometrica,* Vol. 57, pp. 1361-1401.

[37] QUANDT, R. (1960). "Tests of the Hypothesis that a Linear Regression Obeys Two Separate Regimes", *Journal of the American Statistical Association,* Vol. 55, pp. 324-330.

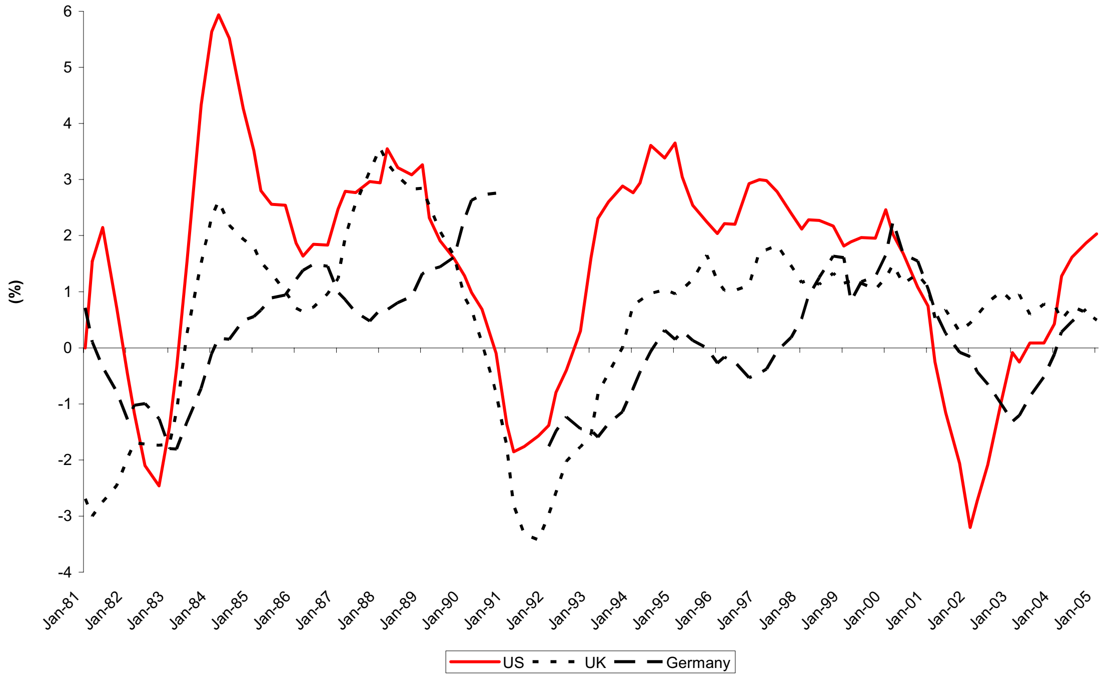
[38] TIMMER, M. P., G. YPMA and B. VAN ARK (2003). *IT in the European Union: Driving Productivity Divergence?,* GGDC Research Memorandum GD-67, Groningen Growth and Development Centre, October.

[39] VAN ARK, B., R. INKLAAR and R. H. MCGUCKIN (2003). "ICT and productivity in Europe and the United States. Where do the differences come from?", *CESifo Economic Studies,* Vol. 49, 3/2003, pp. 295-318.

[40] WASMER, E. (2003). *Interpreting European and US Labour market differences: The specificity of human capital investments,* CEPR Working Paper No. 3780.

[41] WESTERLUND, J. (2006). "Testing for Panel Cointegration with Multiple Structural Breaks", *Oxford Bulletin of Economics and Statistics,* Vol. 68 (1), pp. 101-132.

Figure 1. Employment (annual growth rates)



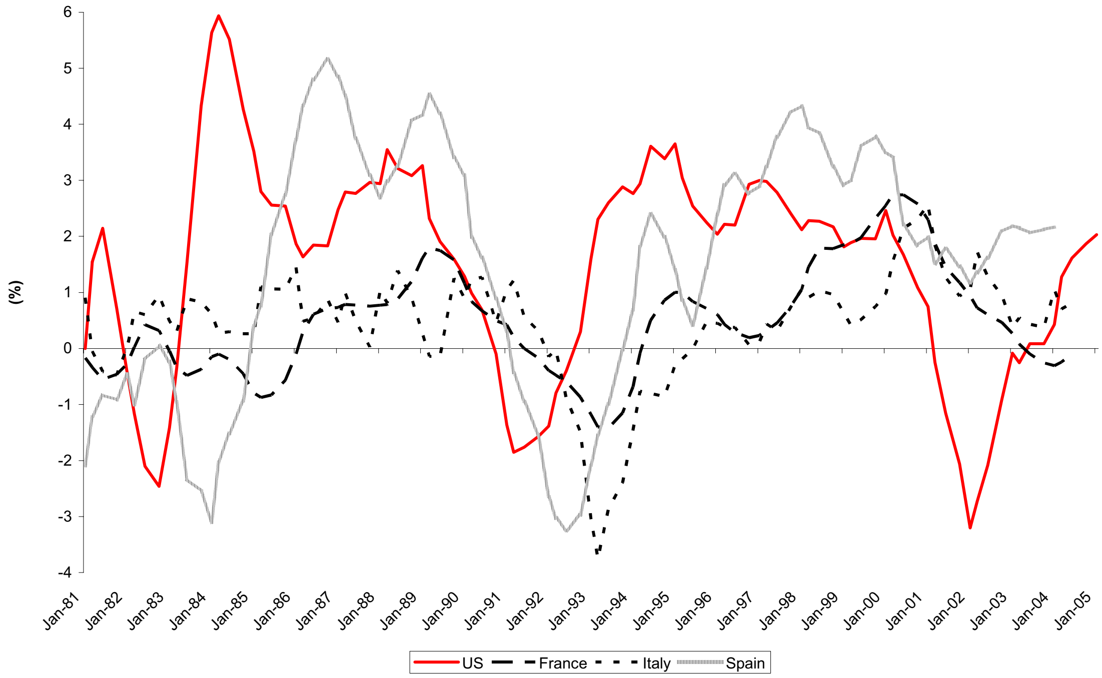


Figure 2. Hours worked (annual growth rates)

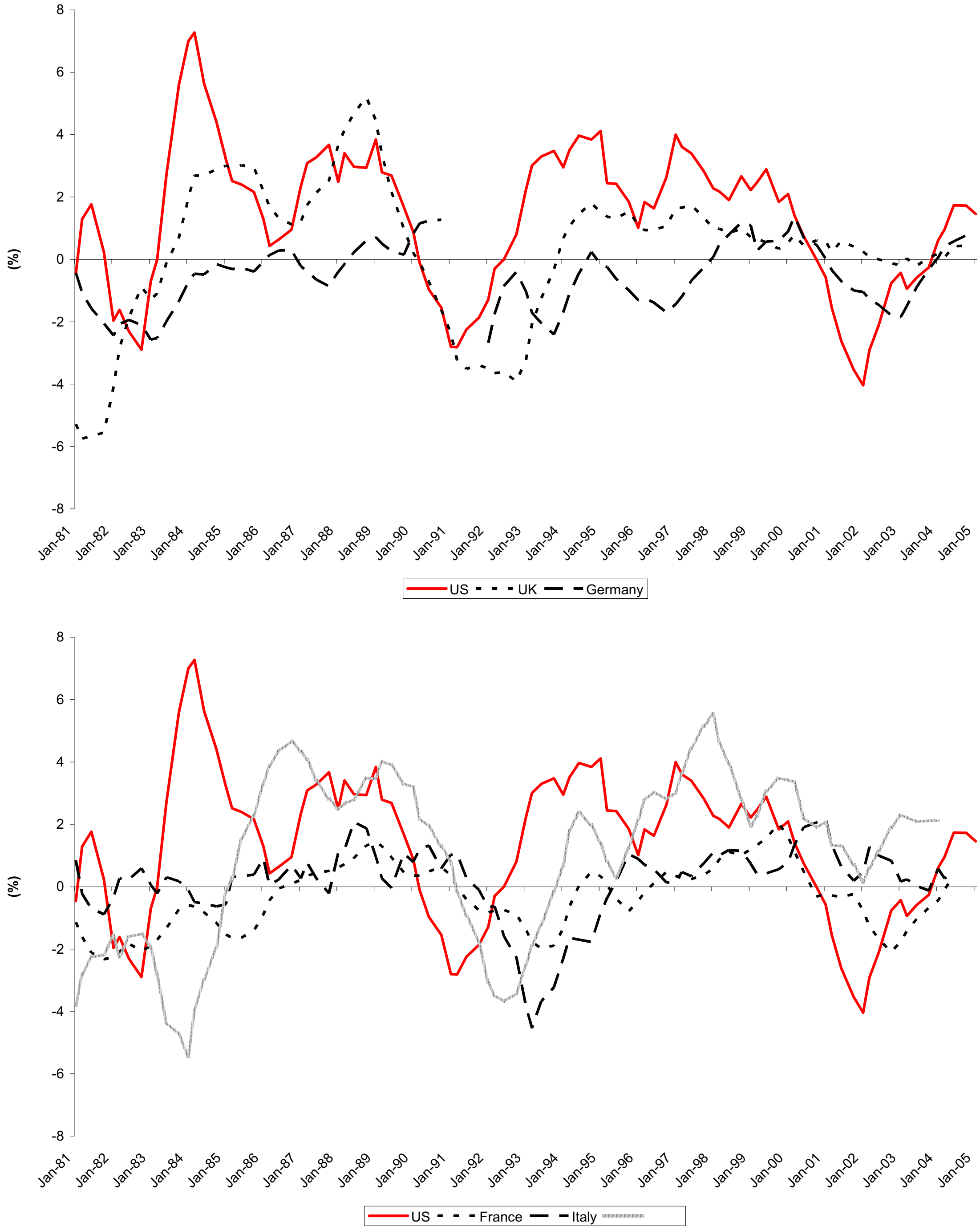
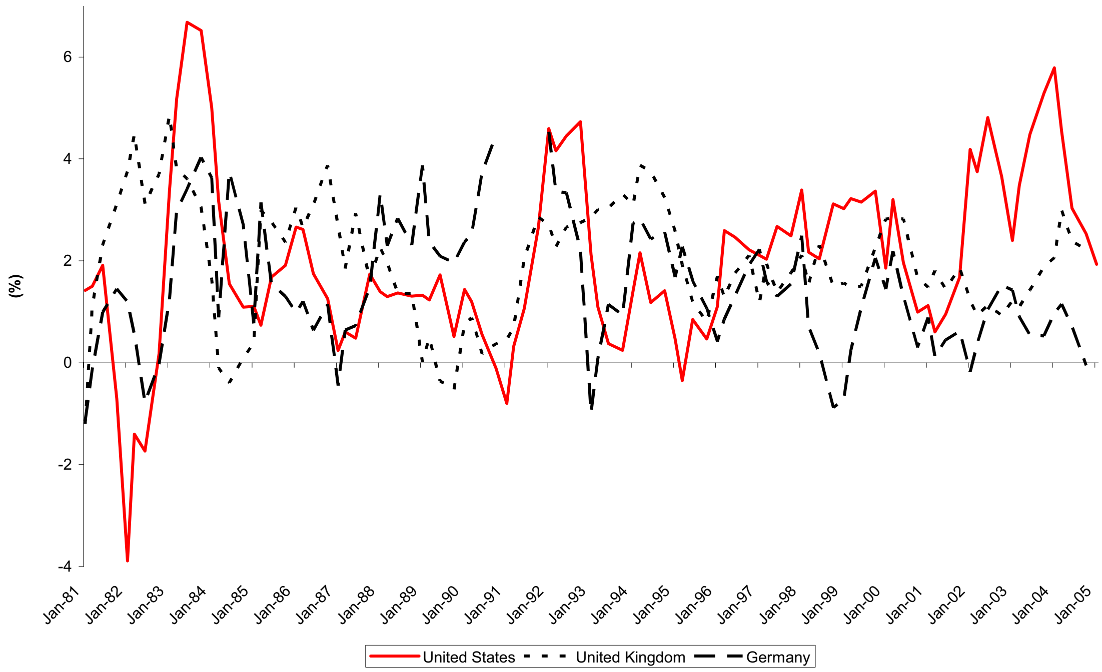


Figure 3a. Real GDP per worker (annual growth rates)



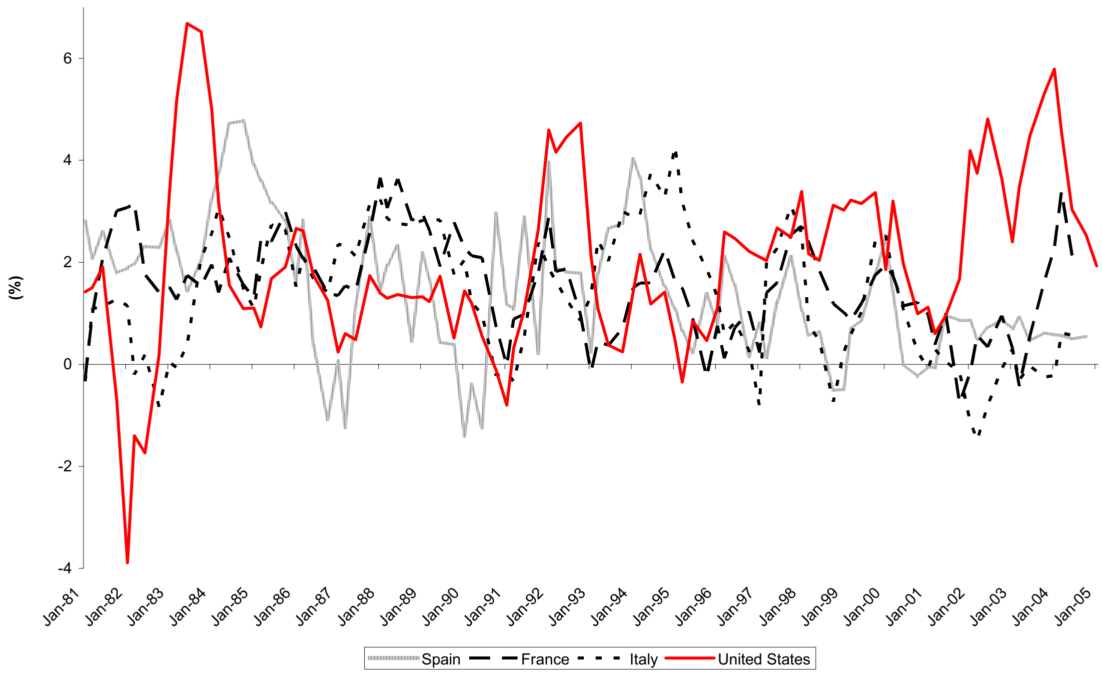
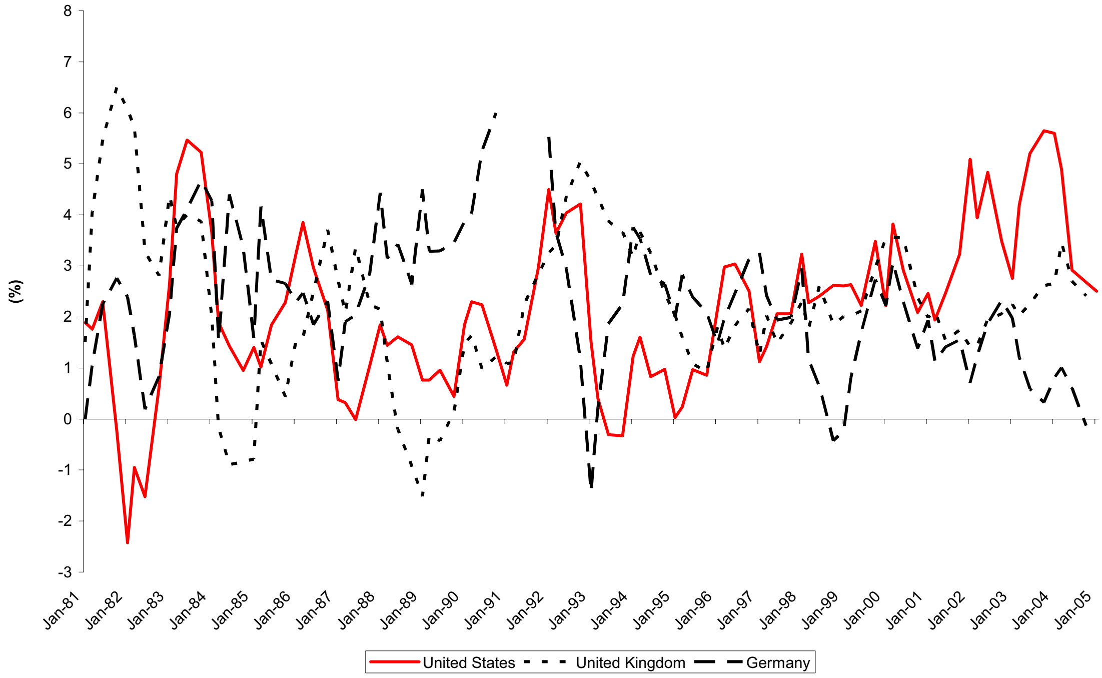


Figure 3b. Real GDP per hour worked (annual growth rates)



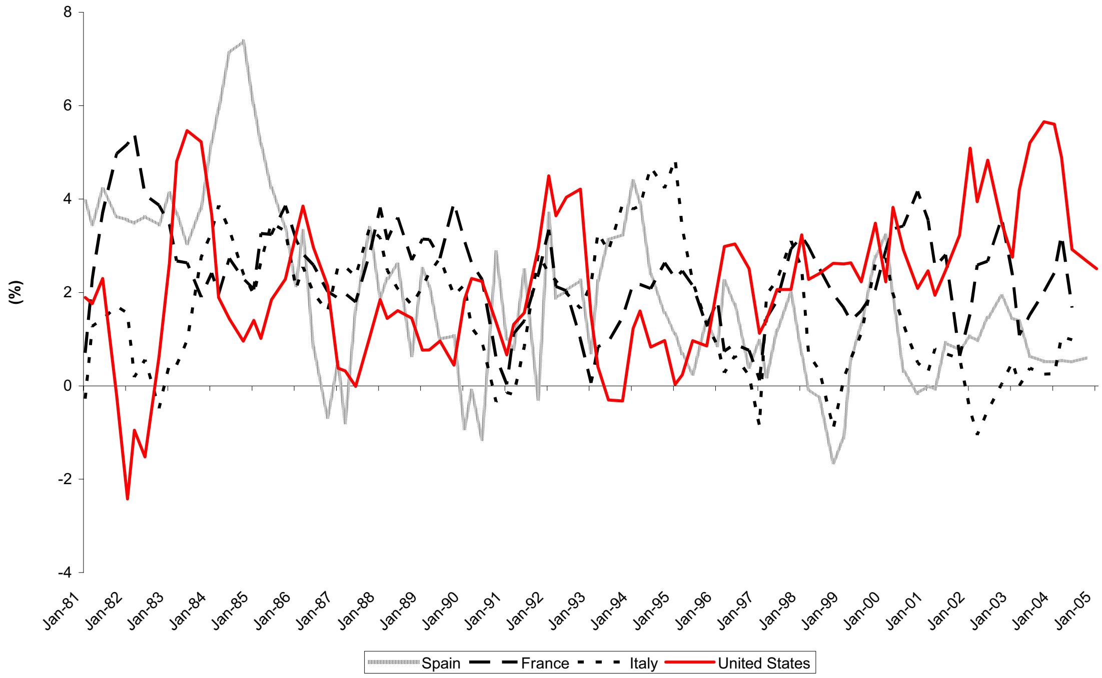
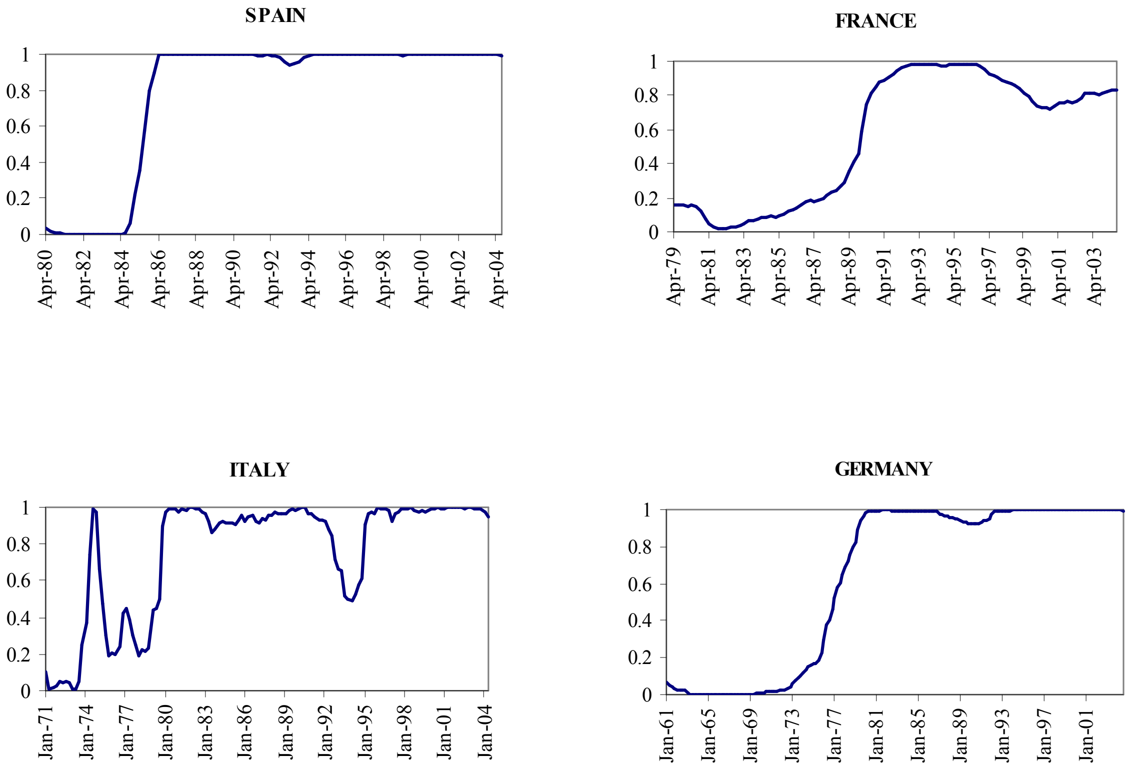


Figure 4
Smoothed probabilities of low mean growth regime
(MS-AR univariate)



Smoothed probabilities of low variance regime
(MS-AR univariate)

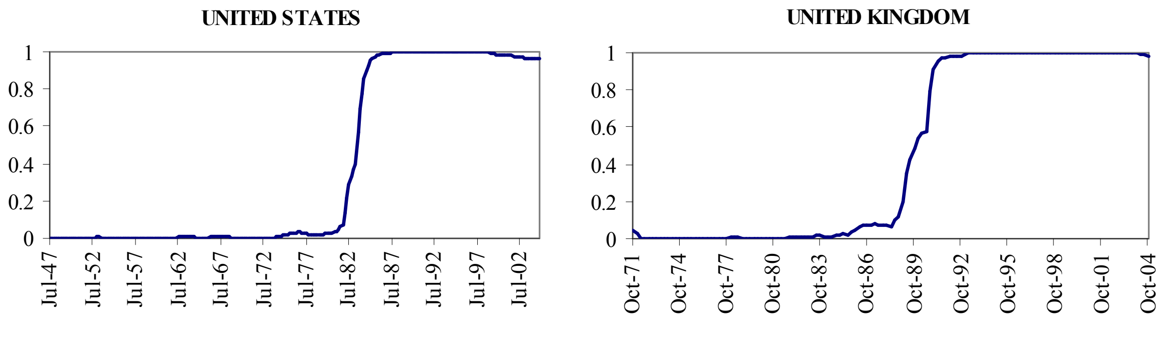


Figure 5
Smoothed and Real Time probabilities of low mean growth regime
(Multivariate Dynamic Factor with MS)

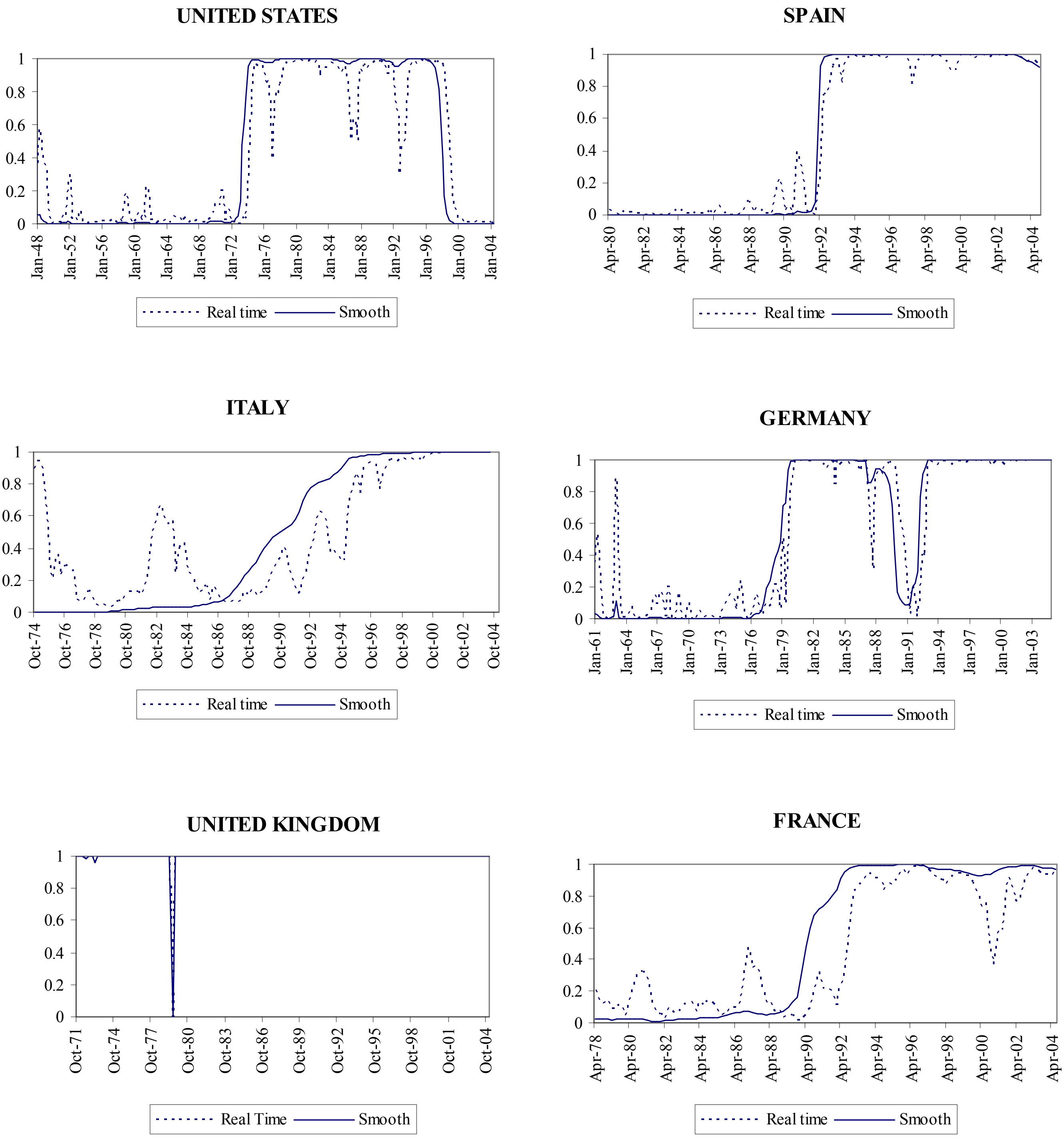


Table 1: Labor productivity, capital-labor ratio and TFP

(Average annual growth rates, %)

	Real GDP per hour worked [1] [2]	Capital-Labor [*] ratio [1] [3]	TFP [4]
United States			
1980-1994	1.6	1.4	0.5
1995-2004	2.8	2.6	1.2
European Union 15			
1980-1994	2.3	2.5	1.1
1995-2004	1.4	1.5	0.7
Germany			
1980-1994	2.7	1.6	1.6
1995-2004	1.6	0.8	1.1
United Kingdom			
1980-1994	2.3	2.3	1.1
1995-2004	2.1	2.5	1.2
France			
1980-1994	2.5	3.3	0.9
1995-2004	2.1	1.4	0.9
Italy			
1980-1994	2.1	3.3	0.7
1995-2004	0.9	1.6	-0.1
Spain			
1980-1994	2.6	3.8	1.9
1995-2004	0.8	1.6	-0.4

[*] Capital refers only to non-residential capital.

Sources:

[1] Banco de España

[2] OECD Productivity Database

[3] AMECO

[4] Updated Database of Timmer, Ypma and van Ark (2003)

Table 2. Estimation of break dates: Summary of results

	Models	AR [1]	AR [2]	Unrestricted Bivariate VAR	Unrestricted Bivariate VAR	Unrestricted trivariate VAR	VECM	VECM
	Variables	P	P	P,R	P,C	P,R,C	2 Cointegration Relations	1 Cointegration Relation
Countries	**Sample period**							
United States	1947.Q1-2005.Q1	1973.Q3 1949.Q2, 1997.Q4	1973.Q2** 1967.Q1 1986.Q2	1973.Q3** 1968.Q3, 1978.Q3	1973.Q3 1951.Q1, 1996.Q1	1973.Q3** 1968.Q2, 1978.Q4	1973.Q3** [3] 1972.Q2, 1974.Q4	-
	1974.Q1-2005.Q1	1996.Q4** 1994.Q3, 1999.Q1	1997.Q1** 1994.Q2 1999.Q3	1997.Q3* 1994.Q1, 2001.Q1	1997.Q4** 1995.Q1, 2000.Q3	1997.Q3 1994.Q2, 2000.Q4	1997.Q3 [3] 1996.Q2, 1998.Q4	-
Spain	1980.Q1-2004.Q4	1986.Q2** 1985.Q2, 1987.Q2	1985.Q4** 1983.Q2 1986.Q3	1986.Q2** 1985.Q2, 1987.Q2	1985.Q3** 1984.Q3, 1986.Q3	1985.Q3** 1984.Q3, 1986.Q3	1990.Q1** 1989.Q1, 1991.Q1	1994.Q1** 1993.Q3, 1994.Q3
	1986.Q1-2004.Q4	-	-	1993.Q3** [4] 1991.Q4, 1995.Q2	2000.Q4 [4] 1995.Q4, >2004.Q4	1993.Q3** [4] 1991.Q4, 1995.Q2	1993.Q4** [4] 1993.Q3, 1994.Q1	1994.Q1** [4] 1993.Q4, 1994.Q2
Italy	1970.Q1-2004.Q3	1980.Q2** 1971.Q1, 1983.Q2	1979.Q4* 1976.Q1 1984.Q4	1980.Q2** 1978.Q3, 1982.Q1	1980.Q2 1978.Q1, 1982.Q3	1980.Q2* 1978.Q3, 1982.Q1	1995.Q3** 1994.Q4, 1996.Q2	1984.Q1** 1983.Q3, 1984.Q3
	1970.Q1-2004.Q3	1995.Q3** 1991.Q2, 1999.Q4	1995.Q1** 1989.Q3 2000.Q3	1995.Q3* 1991.Q4, 1999.Q2	1995.Q3** 1992.Q4, 1998.Q2	1995.Q3** 1993.Q1, 1998.Q1	-	-
France	1978.Q1-2004.Q4	1990.Q3** 1983.Q3, 1997.Q3	1990.Q1** 1984.Q1 1996.Q4	1989.Q3** 1986.Q3, 1992.Q3	1990.Q3 1984.Q1, 1997.Q1	1989.Q3** 1986.Q2, 1992.Q4	1989.Q3* 1986.Q3, 1992.Q3	1988.Q2* 1987.Q1, 1989.Q3
United Kingdom	1971.Q1-2004.Q4	1984.Q2 <1971.Q2, >2004.Q4	1981.Q1 1977.Q1 1992.Q2	1977.Q4 1971.Q3, 1984.Q1	1977.Q4 1973.Q4, 1981.Q4	1977.Q4 1974.Q1, 1981.Q3	1985.Q4** 1985.Q2, 1986.Q2	1985.Q4** 1985.Q1, 1986.Q3
Germany	1960.Q1-2004.Q4	1976.Q4** 1973.Q1, 1980.Q3	1979.Q2** 1977.Q4 1986.Q4	1975.Q2** 1974.Q2, 1976.Q2	1979.Q2** 1977.Q3, 1981.Q1	1977.Q4** 1976.Q4, 1978.Q4	1969.Q3** 1969.Q2, 1969.Q4	-
	1980.Q1-2004.Q4	1992.Q3** 1986.Q4, 1998.Q2	1990.Q3** 1986.Q2 1997.Q3	1992.Q3** 1989.Q1, 1996.Q1	1992.Q3** 1988.Q4, 1996.Q2	1992.Q3* 1988.Q4, 1996.Q2	1993.Q1** 1992.Q4, 1993.Q2	-

Notes: The upper value in each cell corresponds to the estimated breakdate and the two values below are the inferior and superior limits of the 90% confidence interval.

[1] Estimated dates and confidence intervals corresponding to the Supremum Wald statistics sequence.

[2] Estimated dates and confidence intervals corresponding to the Minimum Sum of Square Residuals sequence. To see the statistical relevance in this case we employ a t-test of equality of means.

[3] Imposing unit coefficients in the cointegration relations.

[4] We have re-estimated the VARs starting the sample from 1986 onward, as Spain experimented a severe industrial restructuring process around that time, and perhaps this fact could hide a more recent break.

P: Labor productivity

R: Compensation of employees per hour

C: Private consumption per hour

* Statistical significative at 10% level.

** Statistical significative at 5% level.

Table 3
Test of structural breaks in the quarterly growth of labor productivity

Countries	Test	Change in: Intercept	AR(1) parameter	Joint	Variance
United States	NL [(1)]	0.22	0.26	2.18**	1.85**
	SupW	4.80	3.36	9.46	25.61**
	ExpW	0.75	0.63	2.51	9.66**
Spain	NL [(1)]	1.28**	0.47	2.04**	0.54**
	SupW	32.13**	27.08**	38.29**	19.07**
	ExpW	12.64**	10.28**	15.4**	6.88**
Italy	NL [(1)]	0.67**	0.78**	1.59**	0.75**
	SupW	10.11**	14.3**	16.53**	10.43**
	ExpW	2.64**	4.38**	5.16**	2.52**
France	NL [(1)]	0.40	0.44	0.68	0.11
	SupW	8.39**	4.61	8.40	2.78
	ExpW	2.04**	1.42	2.13	0.33
United Kingdom	NL [(1)]	0.06	0.05	2.57**	2.42**
	SupW	2.24	3.02	5.40	23.77**
	ExpW	0.14	0.32	0.97	10.4**
Germany	NL [(1)]	1.72**	0.06	3.76**	1.66**
	SupW	31.23**	11.54**	43.44**	16.4**
	ExpW	11.82**	2.41**	17.41**	5.84**

Note:
The model used in all tests is an AR(1). Three tests are performed to check the existence of a shift in: i) the intercept, ii) the AR(1) parameter, iii) both of them and iv) the residual variance.
In order to do that we employ three statistics: (1) Nyblom's L (NL), (2) Quandt-Andrews Supremum (SupW) and (3) Andrews-Ploberger Exponential (ExpW).
For the second and third tests we have calculated bootstrap p-values to perform the tests.
[(1)] Joint hypothesis in Nyblom's test considers changes in the intercept, autoregressive parameter and residual variance.
Two asterisks are indicating that we reject the null of no break at 95% confidence level.

Table 4 Tests and estimation of the number of breaks

Pure Structural Change AR(1) model

	SupF [(1)]			Test of unknown number of breaks [(2)]		SupF(k+1/k) [(3)]	
	0 vs 1	0 vs 2	0 vs 3	UDmax	WDmax	2 vs 1	3 vs 2
United States	7.47	12.23*	8.74*	12.23*	13.90*	8.57	3
Spain	40.36*	21.16*	15.20*	40.36*	40.36*	9.4	3.98
Italy	11.73*	10.12*	7.97*	11.73*	11.73*	8	3.78
France	7.7	6.62	6.49	7.7	8.45	4.63	4.35
United Kingdom	5.52	6.63	5.67	7.37	11.11	8.89	8.89
Germany	18.34*	15.59*	11.19*	18.34*	18.34*	9.34	4.69

Partial Structural Change AR(1) model

	SupF [(1)]			Test of unknown number of breaks [(2)]		SupF(k+1/k) [(3)]	
	0 vs 1	0 vs 2	0 vs 3	UDmax	WDmax	2 vs 1	3 vs 2
United States	3.67	8.04*	5.77*	8.04*	9.01*	9.32*	2.08
Spain	36.08*	21.64*	14*	36.08*	36.08*	3.26	2.52
Italy	7.18*	6.03	4.86	7.18	7.18	7.81*	3.84
France	7.74*	7.03*	5.50*	7.74*	7.88	4.7	1.91
United Kingdom	1.34	4.79	3.52	4.79	5.94	6.5	3.29
Germany	14.17*	14.74*	10.80*	14.74*	16.52*	9.83*	1.9

Notes:
One asterisk indicates that the null is rejected at 10% significance level.
[(1)] SupF is a Wald-F statistic that checks the null of no break against the existence of a fixed number of breaks: one (1st column), two (2nd column) or three breaks (3rd column).
[(2)] Double maximum tests to check the null of no break against the alternative of an unknown number of breaks. We have estimated two statistics: UDmax (Unweighted Double Maximum) and WDmax (Weighted Double Maximum).
[(3)] SupF(k+1/k) is a Wald-F statistic that checks the null of the existence of k breaks against the alternative of k+1 breaks (for k = 1, 2).

Table 5 Estimations of Structural Breaks in the mean of quarterly labor productivity growth (1)

	Sup-Wald Statistic			Minimum Least Squares		
Country	Estimation	90% Confidence Interval (2)		Estimation	90% Confidence Interval (2)	
		Min.	Max.		Min.	Max.
USA	1973.Q3	1949.Q2	1997.Q4	1973.Q2	1967.Q1	1986.Q2
	1996.Q4	1994.Q3	1999.Q1	1997.Q1	1994.Q2	1999.Q3
Spain	1986.Q2	1985.Q2	1987.Q2	1985.Q4	1983.Q2	1986.Q3
Italy	1980.Q2	1971.Q1	1983.Q2	1979.Q4	1976.Q1	1984.Q4
	1995.Q3	1991.Q2	1999.Q4	1995.Q1	1989.Q3	2000.Q3
France	1990.Q3	1983.Q3	1997.Q3	1990.Q1	1984.Q1	1996.Q4
UK	1984.Q2	<1971.Q2	>2004.Q4	1981.Q1	1977.Q1	1992.Q2
Germany	1976.Q4	1973.Q1	1980.Q3	1979.Q2	1977.Q4	1986.Q4
	1992.Q3	1986.Q4	1998.Q2	1990.Q3	1986.Q2	1997.Q3

Country	Sub-samples	Mean (% quarterly growth)	Std Dev.	Sub-samples	Mean (% quarterly growth)	Std Dev.
USA	1947.Q1:1973.Q2	0.68	0.09	1947.Q1:1973.Q1	0.68	0.09
	1973.Q2:1996.Q3	0.38	0.09	1973.Q2:1997.Q1	0.36	0.09
	1996.Q4:2005.Q1	1.17	0.18	1997.Q2:2005.Q1	1.29	0.2
Spain	1980.Q1:1986.Q1	0.99	0.13	1980.Q1:1985.Q3	1.05	0.12
	1986.Q2:2004.Q4	0.48	0.11	1985.Q4:2004.Q4	0.45	0.10
Italy	1970.Q1:1980.Q1	0.92	0.23	1970.Q1:1979.Q3	0.97	0.23
	1980.Q2:1995.Q2	0.53	0.07	1979.Q4:1994.Q4	0.53	0.08
	1995.Q3:2004.Q4	0.20	0.09	1995.Q1.2004.Q4	0.20	0.10
France	1978.Q1:1990.Q2	0.72	0.07	1978.Q1:1989.Q4	0.73	0.06
	1990.Q3:2004.Q4	0.47	0.06	1990.Q1:2004.Q4	0.46	0.06
UK	1971.Q2:1984.Q1	0.70	0.16	1971.Q2:1980.Q4	0.60	0.20
	1984.Q2:2004.Q4	0.35	0.05	1981.Q1:2004.Q4	0.34	0.05
Germany	1960.Q4:1976.Q3	1.19	0.14	1960.Q4:1979.Q1	1.16	0.12
	1976.Q4:1992.Q2	0.74	0.09	1979.Q2:1990.Q2	0.64	0.11
	1992.Q3:2004.Q4	0.27	0.06	1990.Q3-2004.Q4	0.27	0.06

(1) All the results apply to a complete or pure structural change AR(1) model.
(2) Estimations of confidence intervals use the asymptotic distribution of the breakdates derived by Bai (1994,1997) based on Picard distribution (Picard, 1985).

Table 6

Structural Break Augmented Fractional Dickey Fuller (SB-AFDF) tests [1]

Quarterly growth of productivity per hour

Model 1: Pure Structural change model

	USA		Spain		Italy		France		UK		Germany	
Order of integration under the null (d)	Statistic	Break Point	Statistic	Break Point	Statistic	Break Point	Statistic	Break Point	Statistic	Break Point	Statistic	Break Point
0.1	-2.464	1978.Q2	-1.362	1985.Q3	-2.679	1993.Q3	-2.702	1992.Q1	-1.777	1984.Q1	-2.687	1990.Q3
0.2	-3.089**	1978.Q2	-1.993	1985.Q3	-2.992	1993.Q3	-2.951	1992.Q1	-2.126	1984.Q1	-3.117**	1990.Q3
0.3	-3.691**	1978.Q2	-2.631	1985.Q3	-3.310**	1993.Q3	-3.391**	1990.Q1	-2.501	1984.Q1	-3.537**	1990.Q3
0.4	-4.267**	1978.Q2	-3.28	1985.Q3	-3.638**	1993.Q3	-3.828**	1990.Q1	-2.90	1984.Q1	-3.959**	1990.Q3
0.6	-5.366**	1978.Q2	-4.635**	1985.Q1	-4.344**	1993.Q3	-4.621**	1990.Q1	-3.781	1984.Q1	-4.841**	1990.Q3
0.7	-5.893**	1978.Q2	-5.359**	1985.Q1	-4.726**	1992.Q4	-4.992**	1990.Q1	-4.256	1984.Q1	-5.310**	1990.Q3
0.8	-6.414**	1978.Q2	-6.088**	1985.Q1	-5.128**	1992.Q4	-5.357**	1990.Q1	-4.754	1984.Q1	-5.800**	1990.Q3
0.9	-6.932**	1978.Q2	-6.822**	1985.Q1	-5.538**	1992.Q4	-5.721**	1990.Q1	-5.269**	1984.Q1	-6.307**	1990.Q3

Model 2: Partial Structural change model

	USA		Spain		Italy		France		UK		Germany	
Order of integration under the null (d)	Statistic	Break Point	Statistic	Break Point	Statistic	Break Point	Statistic	Break Point	Statistic	Break Point	Statistic	Break Point
0.1	-2.14	1996.Q3	-1.751	1987.Q3	-4.044**	1997.Q4	-1.87	1999.Q4	-1.662	1977.Q2	-2.869**	1990.Q3
0.2	-2.35	1996.Q3	-1.948	1987.Q3	-3.794**	1997.Q4	-2.176	1991.Q2	-1.752	1991.Q4	-2.932**	1990.Q3
0.3	-2.847**	1973.Q2	-2.454	1986.Q2	-3.687**	1979.Q4	-2.689	1990.Q1	-1.878	1991.Q4	-3.042**	1990.Q3
0.4	-3.401**	1973.Q2	-2.96	1986.Q2	-4.027**	1979.Q4	-3.262**	1990.Q1	-2.068	1991.Q4	-3.349**	1979.Q1
0.6	-4.528**	1973.Q2	-3.745**	1986.Q2	-4.682**	1979.Q4	-4.241**	1990.Q1	-2.795	1983.Q4	-4.240**	1979.Q1
0.7	-5.107**	1973.Q2	-4.057**	1986.Q2	-5.009**	1979.Q4	-4.685**	1990.Q1	-3.180	1983.Q4	-4.724**	1979.Q1
0.8	-5.693**	1973.Q2	-4.342**	1986.Q2	-5.342**	1979.Q4	-5.115**	1990.Q1	-3.597	1983.Q2	-5.237**	1979.Q1
0.9	-6.281**	1973.Q2	-4.617**	1986.Q2	-5.685**	1979.Q4	-5.537**	1990.Q1	-4.032	1983.Q2	-5.772**	1979.Q1

Notes:

[1] Test developed by Dolado, Gonzalo and Mayoral (2005) to verify the null of I(d) or fractional integration of order d against the alternative of I(0) with deterministic components subject to structural breaks at unknown dates. The break point estimated is the date in which the SB-AFDF (Structural Break Augmented Fractional Dickey Fuller) statistic is minimized.

Model 1 (pure structural change model) has one break in intercept and in the deterministic trend at an unknown date. It corresponds to the crash and changing growth model in Perron's terminology.

Model 2 (partial structural change model) only allows for one break in the intercept at an unknown date. Perron called it crash model.

The number of lags included in each model is the one that minimizes SBIC.

Two asterisks indicate that the null hypothesis is rejected at 95% confidence level.

Table 7: Univariate MS-AR models

	Spain		Italy		France		Germany	
μ_0	0.30	(0.06)	0.35	(0.18)	0.49	(0.08)	0.58	(0.09)
μ_1	1.06	(0.12)	1.15	(0.28)	0.70	(0.08)	1.18	(0.12)
σ^2	0.50	(0.07)	0.45	(0.08)	0.19	(0.08)	1.26	(0.06)
q	0.99	(0.01)	0.97	(0.05)	0.98	(0.04)	0.99	(0.01)
p	0.98	(0.03)	0.90	(0.09)	0.98	(0.04)	0.99	(0.01)
μ_0-μ_1	-0.76		-0.80		-0.21		-0.60	
Prob(St=0)[(1)]	0.67		0.74		0.47		0.44	
Prob(St=1)[(1)]	0.33		0.26		0.53		0.56	
Expected Duration of St=0 (years)[(3)]	22.73		7.35		10.42		27.78	
Expected Duration of St=1 (years)[(3)]	11.36		2.55		11.90		35.71	
% of time spent in St=0	66.67		74.24		46.67		43.75	
Log-likelihood	-109.77		-145.40		-63.07		-272.21	
SBIC	247.05		334.88		167.76		575.40	

	United States		United Kingdom	
μ	0.56	(0.05)	0.59	(0.05)
σ_0^2	1.07	(0.06)	1.16	(0.01)
σ_1^2	0.62	(0.05)	0.38	(0.04)
q	0.995	(0.01)	0.99	(0.01)
p	0.99	(0.01)	0.99	(0.01)
σ_1^2-σ_0^2	-0.45		-0.78	
Prob(St=0)[(2)]	0.55		0.53	
Prob(St=1)[(2)]	0.45		0.47	
Expected Duration of St=0 (years)[(3)]	50.00		31.25	
Expected Duration of St=1 (years)[(3)]	41.67		27.78	
% of time spent in St=0	54.55		52.94	
Log-likelihood	-300.81		-147.06	
SBIC	634.28		323.46	

Notes: Standard deviations in parenthesis.
All models are autoregressive of order 1 except for Italy and France which are of order 4 instead. Estimated autoregressive parameters are not included for brevity reasons but they are available upon request.

(1) Notice that in the top table Prob(St=0) refers to the ergodic probability of staying at low-mean growth state and similarly, Prob(St=1) corresponds to the ergodic probability of being in high-mean growth regime.

(2) In the bottom table Prob(St=0) and Prob(St=1) refer to the ergodic probability of remaining in high-variance and low-variance states, respectively.

(3) The expected duration of each state is the length of time that it is expected to remain in it. It is calculated using the transition probabilities q and p. For instance, the mean duration of St=0 is 1/(1-q).

Table 8
Tests and Estimation of Structural Breaks in a multivariate framework

A. BIVARIATE VAR[1]

Variables: Labor productivity and Compensation of employees per hour.

Country	SupW [2]	ExpW [2]	Breakdate	--90% confidence interval--
United States	19.04	6.87	1973.Q3**	(1968.Q3, 1978.Q3)
	(0.00)	*(0.00)*		
	10.14	2.56	1997.Q3*	(1994.Q1, 2001.Q1)
	(0.09)	*(0.10)*		
Spain	26.6	10.41	1986.Q2**	(1985.Q2, 1987.Q2)
	(0.00)	*(0.00)*		
Italy	12.99	3.47	1980.Q2**	(1978.Q3, 1982.Q1)
	(0.03)	*(0.04)*		
	10.49	2.67	1995.Q3*	(1991.Q4, 1999.Q2)
	(0.08)	*(0.09)*		
France	13.35	3.95	1989.Q3**	(1986.Q3, 1992.Q3)
	(0.02)	*(0.02)*		
United Kingdom	4.87	1.02	1977.Q4	(1971.Q3, 1984.Q1)
	(0.58)	*(0.52)*		
Germany	62.61	27.83	1975.Q2**	(1974.Q2, 1976.Q2)
	(0.00)	*(0.00)*		
	12.14	3.65	1992.Q3**	(1989.Q1, 1996.Q1)
	(0.04)	*(0.03)*		

Variables: Labor productivity and Consumption per hour

Country	SupW [2]	ExpW [2]	Breakdate	--90% confidence interval--
United States	4.65	0.92	1973.Q3	(1951.Q1, 1996.Q1)
	(0.62)	*(0.57)*		
	12.89	4.16	1997.Q4**	(1995.Q1, 2000.Q3)
	(0.03)	*(0.02)*		
Spain	21.1	7.69	1985.Q3**	(1984.Q3, 1986.Q3)
	(0.00)	*(0.00)*		
Italy	9.67	1.99	1980.Q2	(1978.Q1, 1982.Q3)
	(0.11)	*(0.18)*		
	14.99	5.03	1995.Q3**	(1992.Q4, 1998.Q2)
	(0.01)	*(0.01)*		
France	6.75	1.84	1990.Q3	(1984.Q1, 1997.Q1)
	(0.32)	*(0.21)*		
United Kingdom	7.84	1.1	1977.Q4	(1973.Q4, 1981.Q4)
	(0.22)	*(0.47)*		
Germany	40.78	17.18	1979.Q2**	(1977.Q3, 1981.Q1)
	(0.00)	*(0.00)*		
	11.07	3.34	1992.Q3**	(1988.Q4, 1996.Q2)
	(0.06)	*(0.04)*		

B. TRIVARIATE VAR[1]

Variables: Labor productivity, Compensation of employees per hour and Consumption per hour.

Country	SupW [2]	ExpW [2]	Breakdate	--90% confidence interval--
United States	18.73	6.96	1973.Q3**	(1968.Q2, 1978.Q4)
	(0.00)	*(0.00)*		
	11.07	3.18	1997.Q3	(1994.Q2, 2000.Q4)
	(0.15)	*(0.14)*		
Spain	21.71	8.19	1985.Q3**	(1984.Q3, 1986.Q3)
	(0.00)	*(0.00)*		
Italy	12.31	3.07	1980.Q2*	(1978.Q3, 1982.Q1)
	(0.09)	*(0.15)*		
	16.05	5.57	1995.Q3**	(1993.Q1, 1998.Q1)
	(0.02)	*(0.02)*		
France	13.41	4.22	1989.Q3**	(1986.Q2, 1992.Q4)
	(0.06)	*(0.05)*		
United Kingdom	8.41	1.61	1977.Q4	(1974.Q1, 1981.Q3)
	(0.35)	*(0.53)*		
Germany	65.18	29.47	1977.Q4**	(1976.Q4, 1978.Q4)
	(0.00)	*(0.00)*		
	11.61	3.56	1992.Q3*	(1988.Q4, 1996.Q2)
	(0.12)	*(0.10)*		

Notes:
[1] Bivariate and Trivariate VARs are pure structural change models, unrestricted, in first differences and with 1 lag.
[2] SupW and ExpW are the Quandt-Andrews supremum and Andrews-Ploberger exponential transformation of the sequence of F-Wald statistics.
* Statistical significative at 10% level.
** Statistical significative at 5% level.
Values in parenthesis correspond to the asymptotic p-values (calculated with the approximation of Hansen, 1997).

Table 9 Testing for breaks in the common stochastic trend

VECM, with DOLS estimated cointegration vectors

Country	Cointegration Relations	Sample	SupWald	ExpWald	Breakdate	--90% confid. interval---
United States	2	1948.Q4-2004.Q2	8.78	1.95	1973.Q3	(1966.Q3, 1980.Q3)
			(0.31)	*(0.40)*		
	2	1975.Q4-2004.Q2	17.66	5.81	1998.Q1**	(1996.Q4, 1999.Q2)
			(0.01)	*(0.01)*		
United Kingdom	2	1973.Q1-2004.Q4	21.57	6.94	1985.Q4**	(1985.Q2, 1986.Q2)
			(0.00)	*(0.01)*		
	1	1973.Q1-2004.Q4	17.18	4.85	1985.Q4**	(1985.Q1, 1986.Q3)
			(0.01)	*(0.03)*		
Spain	2	1981.Q4-2004.Q4	11.62	3.69	1990.Q1*	(1989.Q1, 1991.Q1)
			(0.12)	*(0.08)*		
	1	1981.Q4-2004.Q4	18.96	5.75	1994.Q1**	(1993.Q3, 1994.Q3)
			(0.01)	*(0.01)*		
	2	1987.Q4-2004.Q4	15.28	4.52	1993.Q4**	(1993.Q3, 1994.Q1)
			(0.03)	*(0.04)*		
	1	1987.Q4-2004.Q4	*20*	*6.82*	1994.Q1**	(1993.Q4, 1994.Q2)
			(0.00)	*(0.01)*		
France	2	1979.Q4-2004.Q3	13.06	4.64	1989.Q3*	(1986.Q3, 1992.Q3)
			(0.07)	*(0.04)*		
	1	1979.Q4-2004.Q3	11.63	3.59	1988.Q2*	(1987.Q1, 1989.Q3)
			(0.12)	*(0.09)*		
Italy	2	1976.Q2-2004.Q3	16.77	5.21	1995.Q3**	(1994.Q4, 1996.Q2)
			(0.01)	*(0.02)*		
	1	1976.Q2-2004.Q3	15.25	4.26	1984.Q1**	(1983.Q3, 1984,Q3)
			(0.03)	*(0.05)*		
Germany	2	1962.Q3-2004.Q4	41.5	16.01	1969.Q3**	(1969.Q2, 1969.Q4)
			(0.00)	*(0.00)*		
	2	1981.Q4-2004.Q4	25.6	8.76	1993.Q1**	(1992.Q4, 1993.Q2)
			(0.00)	*(0.00)*		

VECM, with unit cointegrating coefficients imposed

Country	Cointegration Relations	sample	SupWald	ExpWald	Breakdate	--90% confid. interval---
United States	2	1948.Q4-2004.Q2	27.07	8.98	1973.Q3**	(1972.Q2, 1974.Q4)
			(0.00)	*(0.00)*		
	2	1975.Q4-2004.Q2	10.97	3.03	1997.Q3	(1996.Q2, 1998.Q4)
			(0.15)	*(0.15)*		

Notes: Lags selection based on SBIC criterion for each test.

Variables included in the VECM: Labor Productivity, Compensation per hour and Consumption per hour.

* Statistical significative at 10% level.

** Statistical significative at 5% level.

Table 10 Tests and Estimation of Structural Breaks in a multicountry framework

MULTICOUNTRY (PRODUCTIVITY ONLY)

Country	SupW [1]	ExpW [1]	Breakpoint	--90% confidence interval--
All countries (with US)[2]	16.24 *(0.17)*	6.13 *(0.09)*	1997.Q4*	(1996.Q1, 1999.Q3)
All countries (without US)[2]	15.01 *(0.14)*	5.34 *(0.09)*	1998.Q1*	(1996.Q2, 1999.Q4)
France, Germany, Italy and Spain[2]	23.85 *(0.00)*	9.3 *(0.00)*	1998.Q1**	(1997.Q1, 1999.Q1)
France, Germany and Italy	14.83 *(0.03)*	5.21 *(0.02)*	1992.Q3**	(1989.Q4, 1995.Q2)
France and Germany	9.71 *(0.11)*	2.43 *(0.11)*	1992.Q3	(1987.Q3, 1997.Q3)
Italy and Spain[2]	8.97 *(0.14)*	2.14 *(0.15)*	1995.Q3	(1992.Q2, 1998.Q4)
United Kingdom and United States	8 *(0.21)*	2.4 *(0.12)*	1997.Q3	(1992.Q1, 2003.Q1)

Notes:

(1) SupW and ExpW are the Quandt-Andrews supremum and Andrews-Ploberger exponential transformation of the sequence of F-Wald statistics.

(2) Starting the common sample in 1986, as Spain suffered a hard decrease in productivity in 1985, which dominates all the shifts of the rest of countries and its inclusion or not determines the results of the test.

* Statistical significative at 10% level.

** Statistical significative at 5% level.

Values in parenthesis corresponds to the asymptotic p-values (calculated with the approximation of Hansen, 1997).

All the models refers to complete or pure structural change unrestricted VAR models bivarite, trivariate or multivariate with 1 lag and in first differences.

Table 11: Dynamic Factor Model with Markov Switching

	US		Spain		Italy		Germany		France		UK	
Common permanent component												
γ	0.27	(0.04)	0.22	(0.08)	0.23	(0.08)	0.63	(0.07)	0.22	(0.06)	0.60	(0.06)
μ_0	-0.84	(0.20)	-1.34	(0.83)	-0.36	(0.20)	-0.84	(0.20)	-0.60	(0.30)	-7.57	(1.62)
μ_1	0.61	(0.16)	1.51	(0.96)	0.24	(0.16)	1.09	(0.26)	0.86	(0.41)	6.94	(1.48)
φ	-0.13	(0.16)	-0.21	(0.36)	0.61	(0.17)	-0.42	(0.11)	-0.50	(0.18)	0.27	(0.10)
q_1	0.98	(0.01)	0.99	(0.01)	0.99	(0.01)	0.99	(0.02)	0.99	(0.01)	0.01	(0.01)
p_1	0.99	(0.01)	0.99	(0.01)	0.99	(0.01)	0.97	(0.02)	0.99	(0.02)	0.99	(0.01)
Common transitory component												
λ_1	0.23	(0.05)	-0.22	(0.04)	-0.17	(0.07)	-0.03	(0.09)	-0.19	(0.05)	0.19	(0.10)
λ_2	0.09	(0.03)	-0.13	(0.05)	-0.29	(0.23)	0.06	(0.08)	-0.14	(0.12)	-2.23	(1.27)
λ_3	-0.53	(0.04)	-0.21	(0.04)	-0.37	(0.05)	-0.34	(0.10)	-0.35	(0.07)	0.05	(0.10)
λ_4	0.41	(0.04)	0.23	(0.03)	0.26	(0.04)	0.16	(0.04)	0.05	(0.02)	-0.02	(0.04)
τ	-5.83	(1.12)	-2.03	(0.58)	-	-	-	-	0.10	(0.01)	0.48	(1.45)
φ_{11}*	1.37	(0.06)	1.59	(0.07)	0.91	(0.15)	1.31	(0.17)	1.54	(0.16)	1.43	(0.41)
φ_{12}*	-0.47	(0.04)	-0.63	(0.06)	-0.21	(0.07)	-0.43	(0.11)	-0.55	(0.163)	-0.45	(0.41)
q_2	0.79	(0.13)	0.63	(0.23)	-	-	-	-	0.01	(0.03)	0.99	(0.01)
p_2	0.99	(0.01)	0.97	(0.03)	-	-	-	-	0.99	(0.01)	0.06	(0.69)
Idiosyncratic component												
ψ_{11}	0.91	(0.04)	0.94	(0.05)	0.97	(0.03)	0.97	(0.03)	0.85	(0.07)	0.48	(0.23)
ψ_{21}	1.00	(0.01)	0.94	(0.05)	1.00	(0.01)	0.99	(0.02)	0.99	(0.01)	0.32	(0.56)
ψ_{31}	-0.53	(0.11)	0.97	(0.04)	-0.35	(0.38)	-0.14	(0.15)	-0.42	(0.17)	0.93	(0.04)
ψ_{41}	1.35	(0.06)	-0.76	(0.28)	1.43	(0.10)	1.59	(0.06)	1.66	(0.06)	1.61	(0.06)
ψ_{42}	-0.45	(0.04)	-0.14	(0.10)	-0.51	(0.07)	-0.63	(0.05)	-0.69	(0.05)	-0.65	(0.05)
Regime shifts												
$\gamma(\mu_0\text{-}\mu_1)$	-0.39		-0.62		-0.14		-1.23		-0.32		-8.65	
Prob(S_{1t}=0)(1)	0.30		0.50		0.60		0.67		0.50		0.01	
Prob(S_{1t}=1)(1)	0.70		0.50		0.40		0.33		0.50		0.99	
Prob(S_{2t}=0)(2)	0.03		0.09		-		-		0.01		0.99	
Prob(S_{2t}=1)(2)	0.97		0.91		-		-		0.99		0.01	
Expected duration of S_{1t}=0 (years)(3)	15.63		20.83		25.00		16.67		20.83		0.25	
Expected duration of S_{1t}=1 (years)(3)	35.71		20.83		16.67		8.33		20.83		27.78	
% of time spent in S_{1t}=0	30.43		50.00		60.00		66.67		50.00		0.90	
log likelihood	-143.39		29.74		62.33		-104.08		137.69		-213.27	
SBIC	416.87		50.80		-24.12		316.74		-163.46		544.09	

Notes: Standard deviations in parenthesis.

(1) Prob(S_{1t}=0) refers to the ergodic probability of remaining in low-mean growth state and similarly, Prob(S_{1t}=1) corresponds to the ergodic probability of being in high-mean growth regime for the common permanent component.

(2) Prob(S_{2t}=0) and Prob(S_{2t}=1) are the ergodic probabilities of staying at negative-mean growth or at zero-mean growth states for the common transitory component.

(3) The expected duration of each state is the length of time that it is expected to remain in it. It is calculated with the transition probabilities q_1 and p_1. For instance, the mean duration of S_{1t}=0 is $1/(1-q_1)$.

BANCO DE ESPAÑA PUBLICATIONS

WORKING PAPERS[1]

0527 LUIS J. ÁLVAREZ, PABLO BURRIEL AND IGNACIO HERNANDO: Price setting behaviour in Spain: evidence from micro PPI data.

0528 EMMANUEL DHYNE, LUIS J. ÁLVAREZ, HERVÉ LE BIHAN, GIOVANNI VERONESE, DANIEL DIAS, JOHANNES HOFFMANN, NICOLE JONKER, PATRICK LÜNNEMANN, FABIO RUMLER AND JOUKO VILMUNEN: Price setting in the euro area: some stylized facts from individual consumer price data.

0529 TERESA SASTRE AND JOSÉ LUIS FERNÁNDEZ-SÁNCHEZ: Un modelo empírico de las decisiones de gasto de las familias españolas.

0530 ALFREDO MARTÍN-OLIVER, VICENTE SALAS-FUMÁS AND JESÚS SAURINA: A test of the law of one price in retail banking.

0531 GABRIEL JIMÉNEZ AND JESÚS SAURINA: Credit cycles, credit risk, and prudential regulation.

0532 BEATRIZ DE-BLAS-PÉREZ: Exchange rate dynamics in economies with portfolio rigidities.

0533 ÓSCAR J. ARCE: Reflections on fiscalist divergent price-paths.

0534 M.ª DE LOS LLANOS MATEA AND MIGUEL PÉREZ: Differences in changes in fresh food prices by type of establishment. (The Spanish original of this publication has the same number.)

0535 JOSÉ MANUEL MARQUÉS, FERNANDO NIETO AND ANA DEL RÍO: Una aproximación a los determinantes de la financiación de las sociedades no financieras en España.

0536 S. FABIANI, M. DRUANT, I. HERNANDO, C. KWAPIL, B. LANDAU, C. LOUPIAS, F. MARTINS, T. MATHÄ, R. SABBATINI, H. STAHL AND A. STOKMAN: The pricing behaviour of firms in the euro area: new survey evidence.

0537 LUIS J. ÁLVAREZ AND I. HERNANDO: The price setting behaviour of Spanish firms: evidence from survey data.

0538 JOSÉ MANUEL CAMPA, LINDA S. GOLDBERG AND JOSÉ M. GONZÁLEZ-MÍNGUEZ: Exchange-rate pass-through to import prices in the euro area.

0539 RAQUEL LAGO-GONZÁLEZ AND VICENTE SALAS-FUMÁS: Market power and bank interest rate adjustments.

0540 FERNANDO RESTOY AND ROSA RODRÍGUEZ: Can fundamentals explain cross-country correlations of asset returns?

0541 FRANCISCO ALONSO AND ROBERTO BLANCO: Is the volatility of the EONIA transmitted to longer-term euro money market interest rates?

0542 LUIS J. ÁLVAREZ, EMMANUEL DHYNE, MARCO M. HOEBERICHTS, CLAUDIA KWAPIL, HERVÉ LE BIHAN, PATRICK LÜNNEMANN, FERNANDO MARTINS, ROBERTO SABBATINI, HARALD STAHL, PHILIP VERMEULEN AND JOUKO VILMUNEN: Sticky prices in the euro area: a summary of new micro evidence.

0601 ARTURO GALINDO, ALEJANDRO IZQUIERDO AND JOSÉ MANUEL MONTERO: Real exchange rates, dollarization and industrial employment in Latin America.

0602 JUAN A. ROJAS AND CARLOS URRUTIA: Social security reform with uninsurable income risk and endogenous borrowing constraints.

0603 CRISTINA BARCELÓ: Housing tenure and labour mobility: a comparison across European countries.

0604 FRANCISCO DE CASTRO AND PABLO HERNÁNDEZ DE COS: The economic effects of exogenous fiscal shocks in Spain: a SVAR approach.

0605 RICARDO GIMENO AND CARMEN MARTÍNEZ-CARRASCAL: The interaction between house prices and loans for house purchase. The Spanish case.

0606 JAVIER DELGADO, VICENTE SALAS AND JESÚS SAURINA: The joint size and ownership specialization in banks' lending.

0607 ÓSCAR J. ARCE: Speculative hyperinflations: When can we rule them out?

0608 PALOMA LÓPEZ-GARCÍA AND SERGIO PUENTE: Business demography in Spain: determinants of firm survival.

0609 JUAN AYUSO AND FERNANDO RESTOY: House prices and rents in Spain: Does the discount factor matter?

0610 ÓSCAR J. ARCE AND J. DAVID LÓPEZ-SALIDO: House prices, rents, and interest rates under collateral constraints.

0611 ENRIQUE ALBEROLA AND JOSÉ MANUEL MONTERO: Debt sustainability and procyclical fiscal policies in Latin America.

1. Previously published Working Papers are listed in the Banco de España publications calalogue.

0612 GABRIEL JIMÉNEZ, VICENTE SALAS AND JESÚS SAURINA: Credit market competition, collateral and firms' finance.

0613 ÁNGEL GAVILÁN: Wage inequality, segregation by skill and the price of capital in an assignment model.

0614 DANIEL PÉREZ, VICENTE SALAS AND JESÚS SAURINA: Earnings and capital management in alternative loan loss provision regulatory regimes.

0615 MARIO IZQUIERDO AND AITOR LACUESTA: Wage inequality in Spain: Recent developments.

0616 K. C. FUNG, ALICIA GARCÍA-HERRERO, HITOMI IIZAKA AND ALAN SUI: Hard or soft? Institutional reforms and infraestructure spending as determinants of foreign direct investment in China.

0617 JAVIER DÍAZ-CASSOU, ALICIA GARCÍA-HERRERO AND LUIS MOLINA: What kind of capital flows does the IMF catalyze and when?

0618 SERGIO PUENTE: Dynamic stability in repeated games.

0619 FEDERICO RAVENNA: Vector autoregressions and reduced form representations of DSGE models.

0620 AITOR LACUESTA: Emigration and human capital: Who leaves, who comes back and what difference does it make?

0621 ENRIQUE ALBEROLA AND RODRIGO CÉSAR SALVADO: Banks, remittances and financial deepening in receiving countries. A model.

0622 SONIA RUANO-PARDO AND VICENTE SALAS-FUMÁS: Morosidad de la deuda empresarial bancaria en España, 1992-2003. Modelos de la probabilidad de entrar en mora, del volumen de deuda en mora y del total de deuda bancaria, a partir de datos individuales de empresa.

0623 JUAN AYUSO AND JORGE MARTÍNEZ: Assessing banking competition: an application to the Spanish market for (quality-changing) deposits.

0624 IGNACIO HERNANDO AND MARÍA J. NIETO: Is the Internet delivery channel changing banks' performance? The case of Spanish banks.

0625 JUAN F. JIMENO, ESTHER MORAL AND LORENA SAIZ: Structural breaks in labor productivity growth: The United States vs. the European Union.

BANCO DE ESPAÑA

Unidad de Publicaciones
Alcalá, 522; 28027 Madrid
Telephone +34 91 338 6363. Fax +34 91 338 6488
e-mail: Publicaciones@bde.es
www.bde.es