DOCUMENT DE TREBALL

XREAP2008-6

DECONSTRUCTING SHOCKS AND PERSISTENCE IN OECD REAL EXCHANGE RATES

Syed A. Basher (Qatar Central Bank)
Josep Lluís Carrión-i-Silvestre (AQR-IREA)

XARXA DE REFERÈNCIA EN ECONOMIA APLICADA

Deconstructing Shocks and Persistence in OECD Real Exchange Rates

Syed A. Basher*
Qatar Central Bank

Josep Lluís Carrion-i-Silvestre†
University of Barcelona

May 14, 2008

Abstract

This paper analyzes the persistence of shocks that affect the real exchange rates for a panel of seventeen OECD developed countries during the post-Bretton Woods era. The adoption of a panel data framework allows us to distinguish two different sources of shocks, i.e. the idiosyncratic and the common shocks, each of which may have different persistence patterns on the real exchange rates. We first investigate the stochastic properties of the panel data set using panel stationarity tests that simultaneously consider both the presence of cross-section dependence and multiple structural breaks that have not received much attention in previous persistence analyses. Empirical results indicate that real exchange rates are non-stationary when the analysis does not account for structural breaks, although this conclusion is reversed when they are modeled. Consequently, misspecification errors due to the non-consideration of structural breaks leads to upward biased shocks' persistence measures. The persistence measures for the idiosyncratic and common shocks have been estimated in this paper always turn out to be less than one year.

Keywords: Shock persistence, panel data stationarity tests, multiple structural breaks, cross-section dependence

JEL Classification: C32, C33, E31

*Department of Economic Policies, Qatar Central Bank, P.O. Box 1234, Doha, Qatar. E-mail: bashers@qcb.gov.qa

†Corresponding author: AQR-IREA Research Group, Department of Econometrics, Statistics, and Spanish Economy, University of Barcelona. Av. Diagonal, 690. 08034 Barcelona. Tel: +34 934021826; Fax: +34 93 4021821; E-mail: carrion@ub.edu. J. Ll. Carrion-i-Silvestre acknowledges the financial support of the Ministerio de Ciencia y Tecnología under grant SEJ2005-08646/ECON.

1 Introduction

The debate on the persistence of real exchange rates (RER) has been active for decades. Rogoff (1996) overviews the field and concludes that there seems to exist a consensus in the literature that shocks deviations from the real exchange rates last between 3 to 5 years, deviations that are difficult to explain only from the base of the existence of nominal rigidities. In fact, deviations above one and a half years are considered larger enough as to rule out real rigidities – differentials in productivity or different sectorial economic structures – as a potential determinants of the persistence of these deviations – see Engel and Rogers (2001) and O'Connell and Wei (2002). These controversial assessments are not exempt from criticism if we think of many empirical analyses that have studied the stochastic properties of RER where the main conclusion is that RER can be characterized as non-stationary I(1) stochastic processes. In this case shock persistence is infinite. The investigation on this topic has been recently benefited from the availability of panel data methods, which has increased the amount of empirical investigations. Such panel studies have been used predominantly in testing the long-run purchasing power parity (PPP), which requires that the RER must be stationary so that shocks will only have transitory effects making the RER a mean-reverting stochastic process. However, in this paper we stress the issue that misspecification errors due to either the lack of accounting for cross-section dependence in panel data or the omission of relevant structural breaks can lead to wrong conclusions when aiming at the RER shocks persistence measure.

There are mainly two approaches to assess the stochastic properties of RER in panel data. The first approach tests the null hypothesis of a unit root in the real exchange rates against the alternative hypothesis that PPP holds in the long-run – see e.g. MacDonald (1996), Oh (1996), Wu (1996), and Papell (1997), among others. The second approach considers the null hypothesis of that RER are I(0) stationary processes against the alternative hypothesis that RER are non-stationary I(1) stochastic processes – see e.g. Kuo and Mikkola (2001) and Wagner (2005). Although both approaches are useful, authors such as Taylor (2001) and Bai and Ng (2004a) have argued that, from a conceptual point of view, it is more natural to specify the null hypothesis of stationarity since, in this case, the theory holds under the null hypothesis and rejected when there is strong evidence against it.[1] In this paper, we follow the approach based on panel stationarity tests.

[1] The specification of the null hypothesis of unit root implies that the theory is false, which is contrary to the notion of PPP.

Although the aforementioned studies are based on more powerful techniques and some even provide evidence supporting long-run PPP, they all assume cross-section independence, an assumption that is very unlikely to hold in the context of PPP applications. In fact, the issue of cross-section dependence comes naturally into PPP analysis due to the definition of the base country that is used to define the real exchange rates. O'Connell (1998) first shows the importance of cross-section dependence when assessing the stochastic properties of RER in a panel data framework. The point was further taken up and examined by Lyhagen (2000) and Banerjee, Marcellino and Osbat (2005), who show that panel data unit root statistics tend to conclude in favor of stationarity when cross-section dependence is not considered. This implies that ignoring cross-sectional dependency can lend misleading empirical support to the PPP hypothesis.

Recently Bai and Ng (2004a) propose an approximate factor model that offers a very convenient way to model cross-section dependence in both panel stationary and unit root tests. The factor model that we also apply in our paper has the added advantage that the estimated common factors and idiosyncratic components are consistent whether they are stationary or not. Applying the factor structure to PPP hypothesis, Bai and Ng (2004a) find the presence of one common stationary component on the real exchange rates of fourteen European countries against the dollar. However, the authors are unable to find evidence in favor of PPP due to non-stationarity of the idiosyncratic components.

One important limitation of some proposals in the literature, including the ones cited above, is that the role of the structural breaks is generally ignored when applying either panel unit root or stationarity statistics. It is well known that erroneous omission of structural breaks in the series can lead to deceptive conclusion when performing the unit root tests either in a time series (e.g. Perron, 1989) or in a panel data (e.g. Carrion-i-Silvestre, del Barrio and López-Bazo, 2001) framework. Some recent panel PPP studies have documented the presence of structural breaks in real exchange rates – see e.g. Papell (2002), Im, Lee and Tieslau (2005), and Harris, Leybourne and McCabe (2005). In a motivating paper, Papell (2002) argued that the rise and fall of the dollar in the 1980s may have changed the slope of dollar-based real exchange rates and thereby lend evidence of structural change in the data generating process (DGP). Indeed, Papell's (2002) results provide favorable support for PPP once structural breaks have been accounted for in the computation of panel data unit root tests.

Our analysis simultaneously considers both cross-section dependence and multiple structural breaks, which have not received much attention in previous studies. We tackle the issue of cross-section dependence in three alternative ways including cross-section demeaning, parametric bootstrap methods, and the approximate factor models as in Bai and Ng (2004). To allow for the possibility of structural change in real exchange rates we have utilized the panel stationary tests of Carrion-i-Silvestre, del Barrio and López-Bazo (2005) and Harris et al. (2005), which are flexible enough to account for a large amount of heterogeneity when dealing with multiple structural breaks.

As argued in Papell and Prodan (2006), the traditional interpretation of the PPP hypothesis requires real exchange rate to be stationary in variance around a constant mean in the long-run. Papell and Prodan (2006) tried to reconcile this view in the presence of structural change considering one restricted structural break, so that the long-run mean remains constant. Here we do not follow this approach when considering multiple structural breaks, since we do not impose the restriction that the level of the real exchange rate has to be the same as the one previous to the structural break. Therefore, the evidence showed in this paper has to be seen in terms of whether real exchange rate is stationary in variance once the presence of multiple structural breaks that affect the level of the time series are taken into account. Strictly speaking, the traditional interpretation of the Cassel's (1918) PPP hypothesis would not be fitted in our framework, though the presence of (unrestricted) level shifts in RER has been interpreted in Dornbush and Vogelsang (1991) as evidence in favor of the Balassa-Samuelson notion of PPP.

Finally, we focus on the persistence of the real exchange rate deviations, which is considered as one of the most puzzling empirical regularities in international macroeconomics – see e.g. Rogoff (1996) and Taylor (2001). Our work distinguishes itself from the large literature on this topic by making a distinction between two different sources of shocks that affect the time series. The establishment of this distinction is quite appealing, especially when dealing with RER time series. As noted in O'Connell (1998), RER time series, constructed for instance using the US as the numeraire country, contains two common components, namely, independent variation in the value of the dollar and independent variation in the US price index. These two common components can be interpreted as (maybe part of) common shocks affecting all time series in the panel data. This defines the first source of shocks that can affect the RER time series. However, the deep analysis of persistence should also consider those idiosyncratic shocks that only affect the domestic economy and, hence, whose effects are only restricted to each time series of the panel data set. This defines the second source of shocks for which persistence measures are estimated.

This decomposition, which to the best of our knowledge has not been previously considered in the literature on RER shocks persistence, is appropriate because (nominal and real) exchange rates usually exhibit both high variability within each country over time as well as strong comovements across countries. For instance, European countries often coordinate many of their economic policies, which make exchange rates correlated across countries. This issue is also relevant for policymakers since the symmetry of shocks often play as a candidate for countries to adopt a single monetary union (e.g. European Monetary Union).

Our main results may be summarized as follows. First, we show that the null hypothesis of independence is strongly rejected so that cross-section dependence has to be accounted for when testing the null hypothesis of panel stationary. Second, unreported results show that little evidence is found in favor of the PPP hypothesis when the analysis only considers cross-section dependence but does not account for structural breaks. This conclusion is reversed when cross-section dependence and structural breaks are jointly considered in the computation of the statistics. Thus, the evidence reported in the paper shows the importance of considering the cross-section dependence and the structural breaks when analyzing the PPP hypothesis. Finally, our half-life point estimates are below one year for both the idiosyncratic and the common components. This finding is compatible with the constructed confidence intervals of half-life estimates.

The rest of the paper is organized as follows. In Section 2 we describe the methodology that is applied throughout the paper. Section 3 presents the data set and reports the results of the analysis. Section 4 discusses the measurement of half-life. Finally, Section 5 concludes.

2 Econometric Methodology

2.1 Panel stationarity tests

Hadri (2000) proposes an LM panel data stationarity test without structural breaks, while Carrion-i-Silvestre et al. (2005) extend the analysis to account for the presence of multiple structural breaks. Since the latter proposal encompasses the former one, we proceed to present the approach in Carrion-i-Silvestre et al. (2005). Let $y_{i,t}$ be the DGP for real exchange rates which is given by

$$y_{i,t} = \alpha_i + \sum_{k=1}^{m_i} \theta_{i,k} DU_{i,k,t} + \beta_i t + \sum_{k=1}^{m_i} \gamma_{i,k} DT^*_{i,k,t} + \varepsilon_{i,t} \tag{1}$$

where $t = 1, ..., T$ and $i = 1, ..., N$ indexes the time series and cross-section units, respectively. The dummy variables $DU_{i,k,t}$ and $DT^*_{i,k,t}$ are defined as $DU_{i,k,t} = 1$ for $t > T^i_{b,k}$ and 0 elsewhere; while $DT^*_{i,k,t} = t - T^i_{b,k}$ for $t > T^i_{b,k}$ and 0 elsewhere. The term $T^i_{b,k}$ denotes the k-th date of the break for the i-th time series (individual), $k = 1, ..., m_i, m_i \geq 1$. The parameters α_i and β_i define the time trend, while $\varepsilon_{i,t}$ denotes the disturbance term. Note that the proposal in Hadri (2000) follows from setting $\theta_{i,k} = \gamma_{i,k} = 0$ $\forall i, k$ in (1). The model in (1) includes individual effects, individual structural break effects (i.e. shift in the mean caused by the structural breaks known as temporal effects where $\beta_i \neq 0$) and temporal structural break effects (i.e. shift in the individual time trend where $\gamma_i \neq 0$). In addition, the specification given by (1) considers multiple structural breaks, which are located at different unknown dates and where the number of breaks is allowed to vary across the members of the panel. It is worth mentioning that the case of no structural breaks is embedded in our analysis, since it is possible that there are some individuals that are not affected by the presence of structural breaks. The test statistic is constructed by estimating (1) for every member of the panel and then averaging the N individual stationarity test statistics in Kwiatkowsky, Phillips, Schmidt and Shin (1992) – hereafter, KPSS test $\eta_i(\lambda_i)$. The general expression for the test statistic is

$$LM(\lambda) = N^{-1} \sum_{i=1}^{N} \eta_i(\lambda_i), \tag{2}$$

with $\eta_i(\lambda_i) = \hat{\omega}_i^{-2} T^{-2} \sum_{t=1}^{T} \hat{S}^2_{i,t}$, where $\hat{S}_{i,t} = \sum_{j=1}^{t} \hat{\varepsilon}_{i,j}$ is the partial sum process obtained using the estimated OLS residuals of (1). The term $\hat{\omega}_i^2$ denotes a consistent estimate of the long-run variance of the error $\varepsilon_{i,t}$, which has been estimated following the procedure in Sul, Phillips and Choi (2005) – we use the Quadratic spectral kernel. In (2), λ is defined as the vector $\lambda_i = (\lambda_{i,1}, ..., \lambda_{i,m_i})' = \left(T^i_{b,1}/T, ..., T^i_{b,m_i}/T\right)'$, which indicates the relative position of the dates of the breaks on the entire time period, T, for each individual i. Note that for the test in Hadri (2000) $\lambda_i = 0$ $\forall i$, since there are no structural breaks. Assuming cross-section independence, Hadri (2000) and Carrion-i-Silvestre et al. (2005) show that $LM(\lambda)$ reaches the following sequential limit under the null hypothesis of stationary panel with multiple shifts

$$Z(\lambda) = \frac{\sqrt{N}(LM(\lambda) - \bar{\xi})}{\bar{\varsigma}} \rightarrow N(0,1),$$

where $\bar{\xi}$ and $\bar{\varsigma}$ are the cross-sectional average of the individual mean and variance of $\eta_i(\lambda_i)$, which are defined in Hadri (2000) and Carrion-i-Silvestre et al. (2005).

In order to estimate the number of breaks and their locations, Carrion-i-Silvestre et al. (2005) follow the procedure developed by Bai and Perron (1998), which proceeds in two steps. First, the breakpoints are estimated by globally minimizing the sum of squared residuals for all permissible values of $m_i \leq m^{\max}, i = 1, ..., N$. Second, we use the sequential testing procedure[2] suggested in Bai and Perron (1998) to estimate the number of structural breaks is used. As a result, we obtain the estimation of both the number and position of the structural breaks. This procedure is then repeated N times to obtain the estimated number of breaks and their locations for each individual. Monte Carlo simulations indicate that the test has good size and power in finite sample.

Recently, Harris et al. (2005) have proposed a panel stationarity test statistic without structural breaks. Their specification is based on the following model

$$\begin{aligned} y_{i,t} &= x_{i,t}\beta + z_{i,t} \\ z_{i,t} &= \phi_i z_{i,t-1} + \varepsilon_{i,t}, \end{aligned} \tag{3}$$

where $x_{i,t}$ collects deterministic regressors in a general way – regressors such as a constant, a linear time trend or broken trends. We can obtain the OLS estimated residuals in (3) and, assuming cross-section independence, compute the statistic given by

$$\hat{S}_k = \frac{\hat{C}_k + \hat{c}}{\hat{\omega}\left\{\hat{a}_{k,t}\right\}}, \tag{4}$$

with $\hat{C}_k = T^{-1/2}\sum_{t=k+1}^{T}\hat{a}_{k,t}$ the autocovariance of order k, where $\hat{a}_{k,t} = \sum_{i=1}^{N}\hat{z}_{i,t}\hat{z}_{i,t-k}$, and $\hat{z}_{i,t}$ denotes the OLS residuals in (3). The item $\hat{c} = (T-k)^{-1/2}\sum_{i=1}^{N}\hat{c}_i$, being $\hat{c}_i$ a correction term defined in Harris et al. (2005) and, $\hat{\omega}^2\left\{a_t\right\}$ is a consistent estimate of the long-run variance of

[2]Note that the sequential approach in Bai and Perron (1998) can be used here since under the null hypothesis we have that the units are stationary in variance. Consequently, the consistency on the specification of the number and position of the structural breaks is warranted. Furthermore, the test remains consistent against the alternative hypothesis of I(1) as shown, for instance, in Lee, Huang and Shin (1997), Kurozumi (2002), and Carrion-i-Silvestre (2003).

$\{a_t\}$, which is estimated following the approach in Sul et al. (2005) as above. Under the null hypothesis of stationarity in variance the statistic $\hat{S}_k \to N(0,1)$. In this paper we follow Harris et al. (2005) and use $k = \left[(3T)^{1/2}\right]$.

2.2 Cross-section dependence

The preceding discussion is based on the assumption that time series in the panel data are cross-section independent, which, as argued earlier, is unlikely to hold in the present context. In this paper we account for cross-section dependence in three alternative ways. First, we follow the suggestion in Levin, Lin and Chu (2002) and proceed to remove the cross-section mean. Second, we follow Maddala and Wu (1999) and compute the empirical distribution by means of parametric bootstrap. These two approaches are applied to all test statistics described above. Finally, we apply the factor structure in Bai and Ng (2004) to account for cross-section dependence in the panel. The factor structure is specified for the $\varepsilon_{i,t}$ disturbance term in (1)

$$\varepsilon_{i,t} = F_t'\pi_i + e_{i,t}, \tag{5}$$

where F_t denotes the $(r \times 1)$ vector of common factors, π_i the loadings and $e_{i,t}$ is the idiosyncratic disturbance term. Note that this decomposition permits assessing the stochastic properties of the observed $y_{i,t}$ variables in terms of idiosyncratic and common factor components. The estimation of these components is carried out using principal components method – see Bai and Ng (2004) for further details. Harris et al. (2005) use the same framework as Bai and Ng (2004) to allow cross-section dependence among individuals in the panel data set and propose a test statistic $\left(\hat{S}_k^F\right)$ that tests the null hypothesis of joint variance stationarity of the common and idiosyncratic components – under the null hypothesis $\hat{S}_k^F \to N(0,1)$. Note that the set up in Carrion-i-Silvestre et al. (2005) does not accommodate for common factors to model cross-section dependence.[3]

2.3 Testing for cross-section independence

Recent developments in the literature offer the possibility of testing for the presence of cross-section dependence among individuals. Pesaran (2004) designs a test statistic based on the

[3] Other proposals in the literature that deal with cross-section dependence are O'Connell (1998), who estimates a SUR specification, and Moon and Perron (2004), and Pesaran (2007), who use common factor models as in Bai and Ng (2004a, b).

average of pair-wise Pearson's correlation coefficients $\hat{p}_j$, $j = 1, 2, \ldots, n$, $n = N(N-1)/2$, of the residuals obtained from the estimation of autoregressive (AR) regression models. The CD statistic in Pesaran (2004) is given by

$$CD = \sqrt{\frac{2T}{n}} \sum_{j=1}^{n} \hat{p}_j \to N(0, 1).$$

This statistic tests the null hypothesis of cross-section independence against the alternative hypothesis of dependence.

Besides, Ng (2006) relies on the computation of spacings to test the null hypothesis of independence. In brief, the procedure in Ng (2006) works as follows. First, we get rid of auto-correlation pattern in individual time series through the estimation of an AR model. As in the test in Pesaran (2004), this allows us isolating cross-section dependence from serial correlation. Taking the estimated residuals from the AR regression equations as individual series, we compute the absolute value of Pearson's correlation coefficients ($\bar{p}_j = |\hat{p}_j|$) for all possible pairs of individuals, $j = 1, 2, \ldots, n$, where as above $n = N(N-1)/2$, and sort them in ascending order. As a result, we obtain the sequence of ordered statistics given by $\left\{\bar{p}_{[1:n]}, \bar{p}_{[2:n]}, \ldots, \bar{p}_{[n:n]}\right\}$. Under the null hypothesis that $p_j = 0$ and assuming that individual time series are Normal distributed, $\bar{p}_j$ is half-normally distributed. Furthermore, let us define $\bar{\phi}_j$ as $\Phi\left(\sqrt{T}\bar{p}_{[j:n]}\right)$, where Φ denotes the cdf of the standard Normal distribution, so that $\bar{\phi} = \left(\bar{\phi}_1, \ldots, \bar{\phi}_n\right)$. Finally, let us define the spacings as $\Delta\bar{\phi}_j = \bar{\phi}_j - \bar{\phi}_{j-1}$, $j = 1, \ldots, n$.

Second, Ng (2006) proposes splitting the sample of (ordered) spacings at arbitrary $\vartheta \in (0, 1)$, so that we can define the group of small (S) correlation coefficients and the group of large (L) correlation coefficients. The definition of the partition is carried out through the minimization of the sum of squared residuals

$$Q_n(\vartheta) = \sum_{j=1}^{[\vartheta n]} \left(\Delta\bar{\phi}_j - \bar{\Delta}_S(\vartheta)\right)^2 + \sum_{j=[\vartheta n]+1}^{n} \left(\Delta\bar{\phi}_j - \bar{\Delta}_L(\vartheta)\right)^2,$$

where $\bar{\Delta}_S(\vartheta)$ and $\bar{\Delta}_L(\vartheta)$ denotes the mean of the spacings for each group respectively. Consistent estimate of the break point is obtained as $\hat{\vartheta} = \arg\min_{\vartheta\in(0,1)} Q_n(\vartheta)$, where definition of some trimming is required – we follow Ng (2006) and set trimming at 0.10.

Once the sample has been splitted, we can proceed to test the null hypothesis of cross-section independence in both sub samples. Rejection of the null hypothesis for the small correlations sample will imply rejection for the large correlations sample provided that the statistics are sorted in ascending order. Therefore, the null hypothesis can be tested for the small, large and the whole sample using the Spacing Variance Ratio $SVR(\eta)$ in Ng (2006), with $\hat{\eta} = \left[\hat{\vartheta} n\right]$ being the number of statistics in the small correlations group. Ng (2006) shows that under the null hypothesis that a subset of correlations is jointly zero, the standardized statistic $svr(\eta) \rightarrow N(0,1)$.

One advantage of the approach in Ng (2006) is that it allows us gaining some insight on how pervasive and strong is the cross-section correlation. Thus, if the proportion of correlations in the large correlations group is greater than the one in the small correlations group, then Ng (2006) interprets this feature as evidence of strong correlation, so that the consideration of common factor models to account for the cross-section dependence may represent a good choice. Therefore, the use of these statistics will help us to decide in which panel stationarity statistic we should most base the statistical inference.

3 Empirical Results

We use the same data set used by Pesaran (2007), which consists of quarterly real exchange rates covering the periods 1973Q1 to 1998Q4 ($T = 104$) for 17 OECD countries, namely Australia, Austria, Belgium, Canada, Denmark, Finland, France, Germany, Italy, Japan, Netherlands, New Zealand, Norway, Spain, Sweden, Switzerland, and United Kingdom.[4] The logarithms of the real exchange rates are computed against the U.S. dollar.

Before progressing any further, it is useful to review the case without any structural breaks in the data. To save space we do not present these results, but these are available from the corresponding author on request. As predicted, both Pesaran (2004) and Ng (2006) statistics point to strong presence of cross-section dependence amongst the individuals in the panel data set. We get mixed results after controlling for the cross-section dependence in the data. For instance, when cross-section demeaned data is used, we strongly reject the null hypothesis of variance stationarity using the Hadri (2000) statistics, while it is not rejected when using the $\hat{S}_k$ of Harris et al. (2005). When cross-section dependence is accommodated using the bootstrap distribution, we get favorable results for the PPP hypothesis, regardless of the type of statistics

[4] We are thankful to Takashi Yamagata for making the data available to us. We include the observations for 1973 in our analysis, while Pesaran (2007) starts at 1974.

used. When common factor[5] framework in Bai and Ng (2004a) is used, we are unable to find support for the PPP hypothesis. This result is in accordance with most of previous evidence in the literature cited above,[6] which implies that the PPP hypothesis is not satisfied for the seventeen OECD countries that have been considered in the analysis.

These results, however, are conditional to the assumption that the parameters of the deterministic component of the model are stable throughout the sample period. Consequently, we address the robustness of the above results in the presence of multiple structural breaks using the statistics described in Section 2.

3.1 Panel data stationarity tests with structural breaks

We have estimated the number and position of the structural breaks using the procedure in Bai and Perron (1998) setting $m^{\max} = 5$ as the maximum number of structural breaks – in our analysis this maximum was never attained for any of the time series. The number of break points has been selected with the sequential approach in Bai and Perron (1998) working at the 5% level of significance. The results in Panel A in Table 1 indicate that at least one structural break has been detected for each individual, which shows the potential bias on the estimation of shock persistence measures of those approximations that unattended the structural breaks. The estimated break points are used to compute the statistics in Carrion-i-Silvestre et al. (2005) and Harris et al. (2005).

As a preliminary analysis, we first compute individual stationary statistics to see what fraction of the cross-section units is stationary. Panel A in Table 1 offers the values for the individual $KPSS_i$ and $S_{i,k}$ statistics, as well as the estimated break points. We have also included the simulated critical values at the 10% and 5% level of significance for the individual KPSS statistic. Note that critical values for the $S_{i,k}$ test are not required, since this statistic converges to the standard Normal distribution. Inspection of the individual statistics reveals that the null hypothesis of variance stationarity cannot be rejected in any case for the individual KPSS statistic, while it is rejected in seven cases when using the individual $S_{i,k}$ test at the 5% level of significance. If we combine this information to define panel data statistics and assume that individuals are cross-section independent, we conclude that the null hypothesis of panel

[5] The number of common factors (r) has been estimated using the panel BIC information criterion in Bai and Ng (2002) with up to six common factors.

[6] For instance, Wagner (2005) fails to find support for the PPP hypothesis when applying the Bai and Ng (2004b) methodology to his multi-panel data sets. Similar evidence is also documented in Bai and Ng (2004a).

stationarity cannot be rejected with either version of the $Z(\lambda)$ test, whereas it is rejected if we base on the S_k test – see Panel B in Table 1. However, as discussed above, strong cross-section dependence is evident in the panel, so that we should check if it also present in this case.

Table 2 reports the cross-section correlations based on Pesaran (2004) and Ng (2006). We have estimated an AR regression equation that includes dummy variables to account for the presence of level shifts.[7] As can be seen, the CD statistic strongly rejects the null hypothesis of independence. As for the statistics in Ng (2006), the p-values of the whole $\left(svr^{W}(\hat{\eta})\right)$ and small $\left(svr^{S}(\hat{\eta})\right)$ sample indicate that the null hypothesis of independence cannot be rejected at the 5% level of significance, while the null hypothesis is strongly rejected for the cross-section units in the large group $\left(svr^{L}(\hat{\eta})\right)$. Moreover, the estimated break point $\hat{\eta} = 14$ suggests that the fraction of individuals in the S group is small $\left(\hat{\vartheta} = 0.103\right)$ compared to the correlation coefficients in the L group, which leads us to conclude that strong cross-section correlation among cross-section units is also present when structural breaks are considered in the analysis.

When cross-section dependence is addressed, either through cross-section demeaning or computing the empirical distribution by means of parametric bootstrap[8], all panel data statistics indicate that the null hypothesis of variance stationarity cannot be rejected at the 5% level of significance. See results reported in the columns labeled as 'CS demeaned' and 'bootstrap distribution' in Table 1. This conclusion is reinforced by the $\hat{S}_k^F$ statistic which uses the common factor approach in Bai and Ng (2004a) to model the cross-section dependence. Therefore, we have found strong evidence of stationarity in variance of the real exchange rates for the set of countries that have been considered using both versions of the $Z(\lambda)$ test, and the $\hat{S}_k$ and $\hat{S}_k^F$ statistics when the level of significance is set at the 5% level.[9]

3.2 Structural breaks and real rigidities

Panel A in Table 1 reports the estimated break points obtained from the Bai and Perron (1998) procedure, which are depicted in Figure 1. Except for New Zealand, at least three breaks are found for each country with all breaks occurring during the period 1976Q3 to 1993Q2. From

[7] The AR regression equation in which the statistic is based uses the t-sig criterion in Ng and Perron (1995) to select the order of the autoregressive correction with up to ten lags.

[8] Following Maddala and Wu (1999), we have computed the bootstrap empirical distribution of the statistics using 20,000 replications – we offer the percentiles of interest in Table 1.

[9] Our results are qualitatively different from Harris et al. (2005) who are unable to find favorable support for the PPP hypothesis when applying panel stationary test with cross-section dependence and structural breaks. One possible reason for the discrepancy of research conclusions is the constrained framework imposed in Papell (2002), and maintained in Harris et al. (2005). As mentioned, our results are based on an unconstrained set-up that does not restrict the real exchange rates to return to the levels previous to the structural change.

an historical point of view, the estimated break points are in accordance with events such as oil price shocks, the rise and fall of U.S. dollar in the 1980s and the formation of European Monetary System (EMS). In fact, Papell (2002) identified graphically three major regimes that are likely to have impacted the slopes of real and nominal exchange rates during the post-1973 era. The results in Table 1 reveal that in most cases, the first break occurred during the period 1976Q3 to 1978Q3, which may have resulted due to the oil price shocks in 1973 and 1978. It is possible that events such as the oil embargo or shocks affecting the technological process may change the productivity of cross-section units in different ways, so that differences in productivity can be reduced or increased after the shocks, which may imply a change in the slope of the long-run trend around which the real exchange rates would show stationary fluctuations. In this regard, Alexius (2005) finds that productivity shocks explain about 60–90% of the permanent movements for five out of the six OECD real exchange rates examined.

The second break took place at the beginning of 1980s (between 1981Q1 and 1982Q2), which clearly mimics the start of dollar's appreciation. The third break confirms the transition of dollar's appreciation to depreciation during the period 1986Q2 to 1988Q1. The competitiveness approach emphasizes that real exchange rate depreciations accelerate productivity growth in certain circumstances. For example, a positive demand shock (i.e. real exchange rate depreciation) can increase the measured productivity growth in the tradable goods sector through increased factor utilization, learning-by-doing effects or increasing returns to scale. While these effects are consistent with models of endogenous growth in open economies, there is a related literature arguing the opposite link between cycles and productivity. This strand of the literature identifies reorganization or cleansing recession as reasons that a cyclical downturn could lead to productivity increases – e.g. Saint-Paul (1993). However, empirical results from both sides appear mixed and there is an active empirical debate regarding how permanent the productivity consequences of demand shocks are – see Basu (1998) and Harris (2001) for useful discussion in the productivity-exchange rate debate.

Few countries (mostly European) experienced a fourth break occurring at the beginning of 1990s, which can be explained by the German reunification and/or the formation of EMS. Thus European countries involved in the EMS carried out progressive abolition of any remaining capital controls among the European countries by 1990. In addition, the EMS crisis in September 1992 explains the estimated break points at the beginning of the 1990s. Thus, the exits of Italy and the UK from the exchange rate mechanism of the EMS reflect the detected structural breaks on the fourth and third quarter of 1993 for these countries, respectively. Furthermore, in August 1993 exchange rate bands of the EMS were increased to $\pm 15\%$, which was followed to the adherence of the prospective euro members to the Maastricht conditions on nominal convergence.

As can be seen, the procedure that has been applied in this paper allows the detection of structural breaks that corresponds to some important features that have affected most countries in the panel data set. Furthermore, these estimated break points can be interpreted as the existence of real rigidities that could lead to the estimation of biased measures of shock persistence. These elements were not taken into account in previous studies where testing for the stationarity in variance of the RER using panel data techniques was the main aim.

4 Shock's decomposition and their persistence measurement

In this section we take a fresh look at the "PPP puzzle" using the panel data framework developed above. Briefly, PPP puzzle is the apparent contradiction between the high persistence of shocks to real exchange rate (three to five years) and the high short-term volatility that exhibit (nominal and real) exchange rates. This feature has been investigated in a flurry of papers, where the persistence of the shocks is generally approximated by means of half-life (HL) – the time it takes for 50% of a shock to the real exchange rate to dissipate. Rogoff (1996), while reviewing the empirical literature, reached to the consensus estimate of 3-5 year half-life of real exchange rate adjustment. We are the first to consider the simultaneous presence of cross-section dependence and multiple breaks when focusing on the persistence of the real exchange rate deviation. The application of the factor structure to account for the cross-section dependence allows us to distinguish between common components and country-specific variations in the observed data.

Early panel studies that report shorter half-lives relative to the consensus view were sharply criticized by Murray and Papell (2005) for improper estimation of the autoregressive coefficients that cause downward bias in the half-life. Murray and Papell (2005) address this problem by applying the approximate median-unbiased (MU) estimation method of Andrews and Chen (1994) to post-1973 quarterly data for 20 OECD countries. Their findings suggest an average HL of 3.55 years, which nicely embraces the consensus range. Choi, Mark and Sul (2006) estimate half-lives for deviation from PPP in the range of 5.5 years for a panel of 21 OECD countries during the recent float. However, these studies restrict the autoregressive coefficients to be identical across all cross section units, which is implausible in applied economics since, it implies that each real exchange rate reverts to its respective mean over time at the same rate.

We follow Murray and Papell (2005) to measure the persistence of the shocks, but allow the autoregressive parameters to vary across all cross-section unit. Thus, we have estimated an autoregressive specification for the estimated idiosyncratic disturbance terms and the common factors, which are obtained from the model with multiple structural breaks. Consequently, we compute both idiosyncratic and common HL measures of persistence.

There are different approximations in the literature to estimate the autoregressive model that constitutes the half-life. When time series are thought to be highly persistent and, hence, close to the non-stationary in variance, approaches such as the ones in Pesavento and Rossi (2006) can be employed since these approximations model the time series as a local-to-unity process. When time series are far from the local-to-unity framework, other approaches such as Andrews and Chen (1994), Kilian (1998), and Hansen (1999) can be used. Since, first, the results reported in the previous section reveal that time series are stationary in variance and, second, the initial OLS estimation of the autoregressive models for both components – not reported here to save space – reveal that the largest root of the characteristic polynomial is not close to unity, we have proceeded to compute the HL estimates using the proposal in Andrews and Chen (1994). This procedure also permits to obtain confidence intervals for the autoregressive parameters, and therefore confidence intervals for the idiosyncratic and common HLs can be established as well.

The estimation of the HLs depends on the order of the autoregressive model that is used.[10] Using the method in Andrews and Chen (1994), when we are dealing with an AR(1) model the HL estimate can be directly computed as $HL = \ln(0.5) / \ln(\hat{\alpha}_{MU})$, where $\hat{\alpha}_{MU}$ denotes the median-unbiased autoregressive parameter. However, when the order of the autoregressive model is $p > 1$ the HL estimate has to be obtained from the impulse response function (IRF) that derives from the estimation of the AR(p) model:

$$e_{i,t} = \phi_1 e_{i,t-1} + \cdots + \phi_p e_{i,t-p} + v_{i,t}, \tag{6}$$

where $e_{i,t}$ denotes the idiosyncratic disturbance term in (5). Besides the analysis of the idiosyncratic disturbance term, we may also be interested in investigating the persistence coming from

[10] As above, the selection of the order of the autoregressive model is done using the t-sig information criterion in Ng and Perron (1995) with up to ten lags.

each element of the common factor vector $F_t = (F_{1,t}, \ldots, F_{r,t})'$ in (5), or the effects considering the whole common factor component that affects each individual – i.e. $w_{i,t} = F_t'\pi_i$ in (5) – or the whole stochastic component – i.e. $\varepsilon_{i,t} = F_t'\pi_i + e_{i,t}$ in (5). In these cases, autoregressive models such as the one given in (6) have been estimated replacing $e_{i,t}$ by $F_{j,t}$, $j = 1, \ldots, r$, $w_{i,t}$, or $\varepsilon_{i,t}$, respectively.

Table 3 presents estimates of the autoregressive parameter α and the HL measures that are obtained using the median-unbiased method of Andrews and Chen (1994). In addition to the point estimates, we report the 95% confidence interval for $\hat{\alpha} = \sum_{j=1}^{p} \hat{\phi}_j$, where $\hat{\phi}_j$ denotes the estimated coefficients in (6), and the corresponding HL measures.

Panels A to D in Table 3 report the computations for the idiosyncratic stochastic component ($e_{i,t}$), the stochastic component due to the common factors ($w_{i,t}$), the analysis for each of the common factors ($F_{j,t}, j = 1, \ldots, r$) and the whole stochastic component ($\varepsilon_{i,t}$), respectively. Looking first at the idiosyncratic component (Panel A in Table 3) we find that, although the half-lives vary between countries, all are considerably smaller than one year with a mean of 0.407 years and a median of 0.382 years. For countries such as Sweden and Denmark, real exchange rate variations are apparently dominated by the idiosyncratic components. The common factor estimates (Panel B in Table 3) are very similar but the point estimates of the persistence are very often smaller (with a mean of 0.304 years and a median of 0.265 years) than the one due to the idiosyncratic component. We also see that for most European countries the common shocks are less lasting than for those non-European countries considered in the application, which is consistent with the lower persistence that should be expected when considering integrated economies. If we analyze the effect of each common factor separately we see that their effects are always lower than one year – see Panel C in Table 3. Turning to the estimates for the whole component $\varepsilon_{i,t}$ in Panel D in Table 3, the half-life estimates are roughly in similar magnitude to the previous components, but with a slightly higher mean of 0.421 years and median of 0.396 years. According to these results, real exchange rates' persistence is much smaller than the consensus of 3 to 5 years established by the literature. Note that, save for Austria, in all these situations the confidence intervals are quite narrow and informative when compared to previous estimates in the literature – e.g. Murray and Papell (2005).

To see whether our results are sensitive to the choice of estimation methods, we have also applied the bootstrap method in Kilian (1998) and the grid bootstrap procedure in Hansen (1999) to compute the HLs – these results are not reported to conserve space, although they are available upon request in a companion appendix. We see that the results are quite similar to those based on the median-unbiased estimator. In all cases, the coefficients of the α parameter and the half-life are below unity. Moreover, except for Austria and Denmark when using the Hansen's (1999) method, the confidence interval remains very informative.

Taken together, the above results suggest very rapid adjustments of OECD real exchange rates, faster than the consensus view in the literature. This is interesting and, as suggested in Rogoff (1996), this may indicate that the adjustment to shocks affecting real exchange rate responds to the existence of nominal rather than real market rigidities.[11] In particular, our results emphasize that neglecting structural breaks in the DGP can lead to upward bias in the autoregressive parameters, prompting to conclude that shocks are more persistent than they really are. This is not surprising if we notice that, by definition, structural changes can be seen as few *occasional* shocks having permanent effects on the time series. The methods that usually are employed in the literature to estimate the persistence of shocks investigate the effects of *recurrent* shocks on time series, which is clearly different from occasional shocks. Thus failing to distinguish between these shocks will bias the measures of the shock persistence.

5 Conclusions

In this paper, we test the null hypothesis of stationary of real exchange rate for a panel of seventeen OECD developed countries taking into account both the presence of cross-section dependence and multiple structural breaks. The methodology that we have used is flexible enough to accommodate large degree of heterogeneity with respect to the presence of multiple structural breaks. We have investigated the maintained assumption of cross-section independence in which most previous panel-based PPP studies rely on. Results reveal that strong cross-section correlation is present among individuals and, therefore, a factor structure might help to capture the cross-section dependence. Nevertheless, we have also considered other approaches to account for the presence of cross-section dependence to study the robustness of the conclusions.

Results depend on whether structural breaks are considered. We find evidence in favor of PPP when structural breaks are allowed for, while the evidence breaks down when structural

[11] A recent contribution in this spirit is Imbs et al. (2005) who find relatively less persistence in sub-indices of the CPI than the aggregate CPI. They interpret that this result is due to the underlying heterogeneity in persistence at the level of individual goods and services. Unfortunately Imbs et al. (2005) use first generation panel unit root tests (i.e. those panel data based statistics that assume that time series are cross-sectionally independent), which are highly misleading in the PPP context – see Wagner (2005) for further details.

breaks are omitted. We have also measured the persistence of the idiosyncratic and common shocks to the real exchange rates through the computation of half-life measure. Our half-life point estimates shed light on the PPP puzzle since they turn out to be less than one year for both the idiosyncratic and common factor components used in the analysis. Our results may be interesting in view of recent research that purport to shed light on PPP by exploiting recent advances in panel data econometrics.

References

Alexius, A. (2005). 'Productivity shocks and real exchange rates', *Journal of Monetary Economics*, vol. 52, pp. 555-566.

Andrews, D.W.K. and Chen, H.Y. (1994). 'Approximately median-unbiased estimation of autoregressive models', *Journal of Business and Economic Statistics*, vol. 12, pp. 187-204.

Bai, J. and Perron, P. (1998). 'Estimating and testing linear models with multiple structural changes', *Econometrica*, vol. 66, pp. 47-78.

Bai, J. and Ng, S. (2002). 'Determining the number of factors in approximate factor models', *Econometrica*, vol. 70, pp. 191-221.

Bai, J. and Ng, S. (2004a). 'A new look at panel testing of stationarity and the PPP hypothesis', In *Identification and Inference in Econometric Models: Essays in Honor of Thomas J. Rothenberg*, Don Andrews and James Stock (Ed.), Cambridge University Press.

Bai, J. and Ng, S. (2004b). 'A PANIC attack on unit roots and cointegration', *Econometrica*, vol. 72, pp. 1127-1177.

Banerjee, A., Marcellino, M. and Osbat, C. (2005). 'Testing for PPP: should we use panel methods?', *Empirical Economics*, vol. 30, pp. 77-91.

Basu, S. (1998). 'Technology and business cycles: How well do standard models explain the facts?', In *Beyond Shocks: What Causes Business Cycles?* Federal Reserve Bank of Boston Conference Series No. 42, pp. 207-255.

Caner, M. and Kilian, L. (2001). 'Size distortions of tests of the null hypothesis of stationarity: evidence and implications for the PPP debate', *Journal of International Money and Finance*, vol. 20, pp. 639-657.

Carrion-i-Silvestre, J.L. (2003). 'Breaking date misspecification error for the level shift KPSS test', *Economics Letters*, vol. 81, pp. 365-371.

Carrion-i-Silvestre, J.L., Del Barrio-Castro, T. and López-Bazo, E. (2001). 'Level shifts in a panel data based unit root test. An application to the rate of unemployment', *Proceedings of the 2001 North American Econometric Society.*

Carrion-i-Silvestre, J.L., Del Barrio-Castro, T. and López-Bazo, E. (2005). 'Breaking the panels: an application to the GDP per capita', *Econometrics Journal*, vol. 8, pp. 159-175.

Cassel, G. (1918). 'Abnormal deviation in international exchanges', *Economic Journal*, vol. 28, pp. 413-415.

Choi, C-Y., Mark, N. and Sul, D. (2006). 'Unbiased estimation of the half-life to PPP convergence in panel data', *Journal of Money, Credit and Banking*, vol. 38, pp. 921-938.

Dornbusch, R. and Vogelsang, T. (1991). 'Real exchange rates and purchasing power parity', In *Trade Theory and Economic Reform: North, South and East, Essays in Honor of Bela Balassa*, Cambridge, MA, Basil Blackwell.

Engel, C. and Rogers, J.H. (2001). 'Deviations from purchasing power parity: causes and welfare costs', *Journal of International Economics*, vol. 55, pp. 29-57.

Hadri, K. (2000). 'Testing for stationarity in heterogeneous panel data', *Econometrics Journal*, vol. 3, pp. 148-61.

Hansen, B.E. (1999). 'The grid bootstrap and the autoregressive model', *Review of Economics and Statistics*, vol. 81, pp. 594-607.

Harris, R. (2001). 'Is there a case for exchange rate induced productivity change?', *Discussion Paper No. 0110*, Centre for International Economic Studies.

Harris, D., Leybourne, S., and McCabe, B. (2005). 'Panel stationarity tests for purchasing power parity with cross-sectional dependence', *Journal of Business and Economic Statistics*, vol. 23, pp. 395-409.

Hegwood, N.D. and Papell, D.H. (1998). 'Quasi purchasing power parity', *International Journal of Finance and Economics*, vol. 3, pp. 279-289.

Im, K-S., Lee, J. and Tieslau, M. (2005). 'Panel LM unit-root tests with level shifts', *Oxford Bulletin of Economics and Statistics*, vol. 67, pp. 393-419.

Imbs, J., Mumtaz, H., Ravn, M.O. and Rey, H. (2005). 'PPP strikes back: aggregation and the real exchange rate', *Quarterly Journal of Economics*, vol. 120, pp. 1-43.

Kilian, L. (1998). 'Small-sample confidence intervals for impulse response functions', *Review of Economics and Statistics*, vol. 80, pp. 218-230.

Kuo, B.-S. and Mikkola, A. (2001). 'How sure are we about purchasing power parity? Panel evidence with the null of stationary real exchange rates', *Journal of Money, Credit, and Banking*, vol. 33, pp. 767-789.

Kurozumi, E. (2002). 'Testing for stationarity with a break', *Journal of Econometrics*, vol. 108, pp. 63-99.

Kwiatkowski, D., Phillips, P.C.B., Schmidt, P.J. and Shin, Y. (1992). 'Testing the null hypothesis of stationarity against the alternative of a unit root: how sure are we that economic time series have a unit root', *Journal of Econometrics*, vol. 54, pp. 159-78.

Lee, J., Huang, C.J. and Shin, Y. (1997). 'On stationary tests in the presence of structural breaks', *Economics Letters*, vol. 55, pp. 165-172.

Levin, A., Lin, C.F. and Chu, J. (2002). 'Unit root tests in panel data: asymptotic and finite-sample properties', *Journal of Econometrics*, vol. 108, pp. 1-24.

Lyhagen, J. (2000). 'Why not use standard panel unit root test for testing PPP', *Working Paper Series in Economics and Finance* 413, Stockholm School of Economics.

MacDonald, R. (1996). ‘Panel unit root tests and real exchange rates’, *Economics Letters*, vol. 50, pp. 7-11.

Maddala, G.S. and Wu, S. (1999). ‘A comparative study of unit root tests with panel data and a new simple test’, *Oxford Bulletin of Economics and Statistics*, vol. 61, pp. 631-52.

Moon, R.H. and Perron, B. (2004). ‘Testing for unit root in panels with dynamic factors’, *Journal of Econometrics*, vol. 122, pp. 81-126.

Murray, C.J. and Papell, D.H. (2005). ‘Do panels help solve the purchasing power parity puzzle?’, *Journal of Business and Economic Statistics*, vol. 23, pp. 410-415.

Ng, S. and Perron, P. (1995). ‘Unit root tests in ARMA models with data dependent methods for the selection of the truncation lag’, *Journal of the American Statistical Association*, vol. 90, pp. 268-281.

Ng, S. (2006). ‘Testing cross-section correlation in panel data using spacings’, *Journal of Business and Economic Statistics*, vol. 24, pp. 12-23.

O’Connell, P.G.J. (1998). ‘The overvaluation of the purchasing power parity’, *Journal of International Economics*, vol. 44, pp. 1-19.

O’Connell, P.G.J. and Wei, S.J. (2002). ‘The bigger they are, the harder they fall: retail price differences across U.S. cities’, *Journal of International Economics*, vol. 56, pp. 21-53.

Oh, K-Y. (1996). ‘Purchasing power parity and unit root tests using panel data’, *Journal of International Money and Finance*, vol. 15, pp. 405-418.

Papell, D. (1997). ‘Searching for stationarity: purchasing power parity under the current float’, *Journal of International Economics*, vol. 43, pp. 313-332.

Papell, D. (2002). ‘The great appreciation, the great depreciation, and the purchasing power parity hypothesis’, *Journal of International Economics*, vol. 57, pp. 51-82.

Papell, D. and Prodan, R. (2006). ‘Additional evidence of long-run purchasing power parity with restricted structural change’, *Journal of Money, Credit and Banking*, vol. 38, pp. 1329-1350.

Perron, P. (1989). ‘The great crash, the oil price shock, and the unit root hypothesis’, *Econometrica*, vol. 57, pp. 1361-1401.

Pesaran, M.H. (2004). ‘General diagnostic tests for cross section dependence in panels’, *Cambridge Working Papers in Economics*, No. 435, University of Cambridge.

Pesaran, M.H. (2007). ‘A simple panel unit root test in the presence of cross section dependence’, *Journal of Applied Econometrics*, vol. 22, pp. 265-312.

Pesavento, E. and Rossi, B. (2006). ‘Small sample confidence intervals for multivariate IRFs at long horizons’, *Journal of Applied Econometrics*, vol. 21, pp. 1135-1155.

Rogoff, K. (1996). ‘The purchasing power parity’, *Journal of Economic Literature*, vol. 34, pp. 647-668.

Saint-Paul, G. (1993). ‘Productivity growth and the structure of the business cycle’, *European Economic Review*, vol. 37, pp. 861-83.

Sul D., Phillips P.C.B. and Choi C.Y. (2005). ‘Prewhitening bias in HAC estimation’, *Oxford Bulletin of Economics and Statistics*, vol. 67, pp. 517-546.

Taylor, A.M. (2001). ‘Potential pitfalls for the purchasing-power-parity puzzle? Sampling and specification biases in mean-reversion tests of the law of one price’, *Econometrica*, vol. 69, pp. 473-498.

Wagner, M. (2005). ‘On PPP, unit roots and panels’, *Economics Series* 176, Institute for Advanced Studies.

Wu, Y. (1996). ‘Are real exchange rates nonstationary? Evidence from a panel-data test’, *Journal of Money, Credit and Banking*, vol. 28, pp. 54-63.

Table 1: Individual and panel data stationarity tests with multiple structural breaks

Panel A: Individual statistics

		Critical values						
	$KPSS_i$	10%	5%	$S_{i,k}$	$T^i_{b,1}$	$T^i_{b,2}$	$T^i_{b,3}$	$T^i_{b,4}$
Australia	0.0213	0.059	0.069	-0.284	1976Q4	1984Q2	1988Q1	1992Q2
Austria	0.0305	0.100	0.126	2.326	1977Q2	1981Q1	1986Q2	
Belgium	0.0362	0.063	0.076	1.838	1976Q3	1981Q1	1986Q2	1990Q1
Canada	0.0314	0.054	0.062	0.843	1978Q2	1984Q1	1987Q4	1993Q2
Denmark	0.0228	0.100	0.126	3.051	1977Q2	1981Q1	1986Q2	
Finland	0.0365	0.076	0.091	1.017	1982Q2	1986Q4	1992Q3	
France	0.0283	0.099	0.124	1.391	1977Q2	1981Q2	1986Q2	
Germany	0.0369	0.062	0.074	2.341	1977Q1	1981Q1	1986Q2	1990Q2
Italy	0.0373	0.071	0.083	0.654	1981Q1	1986Q2	1992Q4	
Japan	0.0422	0.100	0.128	-0.711	1977Q1	1981Q2	1986Q1	
Netherlands	0.0372	0.101	0.127	1.987	1976Q3	1981Q1	1986Q2	
New Zealand	0.0163	0.117	0.143	0.127	1983Q1	1986Q4		
Norway	0.0283	0.055	0.062	0.706	1976Q3	1982Q2	1986Q4	1992Q4
Spain	0.0248	0.056	0.065	0.708	1978Q2	1982Q1	1986Q2	1993Q2
Sweden	0.0442	0.055	0.063	1.703	1976Q3	1981Q4	1986Q4	1992Q4
Switzerland	0.0253	0.099	0.125	2.053	1977Q2	1981Q1	1986Q2	
UK	0.0406	0.055	0.063	0.153	1978Q3	1982Q2	1987Q1	1992Q3

Panel B: Panel Stationarity Tests

	Independence		CS demeaned		Bootstrap distribution			
	Test	p-val	Test	p-val	90%	95%	97.5%	99%
$Z(\lambda)$ Hom.	-2.237	0.987	-1.689	0.954	6.780	7.888	8.874	10.279
$Z(\lambda)$ Het.	-1.983	0.976	-1.921	0.973	6.313	7.464	8.548	10.002
$\hat{S}_k$	1.965	0.025	-0.253	0.600	1.639	1.974	2.264	2.575

Panel Stationarity test with common factors

	Test	p-val	$\hat{r}$	$\hat{r}_1$
$\hat{S}^F_k$	1.519	0.064	6	-

Notes: $KPSS_i$ and $S_{i,k}$ report individual test statistic in Kwiatkowsky et al. (1992) and Harris et al. (2005), respectively. $Z(\lambda)$ Hom. and $Z(\lambda)$ Het. report the panel statistics in Carrion-i-Silvestre et al. (2005), while $\hat{S}_k$ represents the panel stationary statistics in Harris et al. (2005). The column CS demeaned refers to the cross-section demeaned procedure in Levin et al. (2002), while the column independence refers the case where cross-section units are independent. $T^i_{b,k}$ denote the k-th date of the break for i-th individual. $\hat{r}$ denotes estimated number of common factors.

Table 2: Cross-section correlation tests

	Test statistic and p-value
CD Statistic	64.586 (0.000)
$\hat{\eta}$	14
$svr^{W}\ (\hat{\eta})$	-1.147 (0.874)
$svr^{L}\ (\hat{\eta})$	6.423 (0.000)
$svr^{S}\ (\hat{\eta})$	-0.587 (0.722)

Notes: All the statistics in the table specify the null hypothesis of cross-section independence. CD refers to the test statistic proposed in Pesaran (2004). $svr^{W}(\hat{\eta})$, $svr^{L}(\hat{\eta})$, and $svr^{S}(\hat{\eta})$ report Ng's (2006) spacing variance ratio for the whole, large, and small sample, respectively. P-values between parentheses. $\hat{\eta}$ indicates the estimated break point in the sample. We use the $t - sig$ criterion in Ng and Perron (1995) to select the order of the autoregressive correction with up to ten lags.

Table 3: Median-unbiased-based autoregressive parameter and Half-life (in years) estimates for the idiosyncratic and common components

Panel A: Idiosyncratic component

		$\hat{\alpha}_{MU}$ parameter			HL estimates		
			95% CI			95% CI	
	Lags	Point	Lower	Upper	Point	Lower	Upper
Australia	0	0.625	0.457	0.819	0.369	0.221	0.868
Austria	3	0.702	0.530	0.853	0.484	0.275	1.003
Belgium	1	0.480	0.310	0.660	0.241	0.181	0.388
Canada	0	0.711	0.521	0.856	0.508	0.266	1.115
Denmark	8	0.775	0.524	0.931	0.703	0.271	3.187
Finland	3	0.709	0.527	0.851	0.501	0.273	1.006
France	3	0.647	0.455	0.791	0.370	0.229	0.685
Germany	3	0.653	0.475	0.806	0.382	0.238	0.744
Italy	4	0.329	0.022	0.531	0.186	0.128	0.273
Japan	0	0.699	0.556	0.880	0.484	0.295	1.356
Netherlands	0	0.722	0.527	0.858	0.532	0.271	1.132
New Zealand	0	0.591	0.372	0.754	0.330	0.175	0.614
Norway	7	0.300	-0.040	0.508	0.179	0.120	0.257
Spain	1	0.735	0.593	0.870	0.499	0.320	1.138
Sweden	0	0.738	0.564	0.878	0.571	0.303	1.332
Switzerland	3	0.569	0.361	0.734	0.305	0.196	0.534
UK	0	0.541	0.240	0.674	0.282	0.122	0.439

Panel B: Common factor component

		$\hat{\alpha}_{MU}$ parameter			HL estimates		
			95% CI			95% CI	
	Lags	Point	Lower	Upper	Point	Lower	Upper
Australia	3	0.628	0.417	0.780	0.347	0.214	0.662
Austria	4	0.523	0.249	0.688	0.265	0.166	0.424
Belgium	4	0.516	0.284	0.684	0.260	0.175	0.417
Canada	7	0.353	0.055	0.506	0.193	0.132	4.784
Denmark	4	0.538	0.281	0.706	0.275	0.174	0.447
Finland	3	0.552	0.331	0.721	0.292	0.187	0.517
France	4	0.493	0.226	0.668	0.246	0.162	0.399
Germany	4	0.520	0.276	0.685	0.263	0.173	0.419
Italy	4	0.347	0.029	0.553	0.192	0.129	0.297
Japan	1	0.699	0.537	0.850	0.442	0.275	0.981
Netherlands	4	0.509	0.257	0.682	0.256	0.168	0.415
New Zealand	5	0.734	0.523	0.896	0.538	0.268	1.433
Norway	3	0.535	0.308	0.691	0.277	0.181	0.456
Spain	4	0.382	0.089	0.571	0.202	0.137	0.310
Sweden	0	0.700	0.530	0.869	0.486	0.273	1.230
Switzerland	1	0.667	0.496	0.817	0.391	0.248	0.748
UK	3	0.492	0.241	0.660	0.246	0.165	0.383

Table 3 (Cont.): Autoregressive parameter and Half-life (in years) estimates for the idiosyncratic, common and whole stochastic components

Panel C: Analysis of common factors

		$\hat{\alpha}_{MU}$ parameter			HL estimates		
			95% CI			95% CI	
Factor	Lags	Point	Lower	Upper	Point	Lower	Upper
1	4	0.485	0.239	0.662	0.243	0.164	0.396
2	3	0.637	0.485	0.762	0.361	0.243	0.559
3	1	0.682	0.531	0.825	0.398	0.269	0.743
4	0	0.702	0.533	0.859	0.490	0.275	1.142
5	0	0.689	0.586	0.896	0.466	0.325	1.573
6	4	0.366	0.050	0.574	0.197	0.132	0.320

Panel D: Whole stochastic component

		$\hat{\alpha}_{MU}$ parameter			HL estimates		
			95% CI			95% CI	
	Lags	Point	Lower	Upper	Point	Lower	Upper
Australia	0	0.613	0.412	0.785	0.354	0.196	0.716
Austria	5	0.755	0.500	0.973	0.670	0.250	12.555
Belgium	1	0.594	0.409	0.764	0.327	0.212	0.611
Canada	0	0.725	0.542	0.873	0.539	0.283	1.276
Denmark	5	0.686	0.402	0.913	0.484	0.209	3.482
Finland	3	0.683	0.521	0.821	0.452	0.267	0.790
France	5	0.679	0.473	0.818	0.382	0.237	0.692
Germany	3	0.720	0.552	0.882	0.524	0.294	1.527
Italy	4	0.084	-0.266	0.326	0.137	0.099	0.185
Japan	0	0.730	0.552	0.884	0.550	0.292	1.406
Netherlands	4	0.640	0.440	0.786	0.391	0.223	0.711
New Zealand	3	0.657	0.494	0.792	0.396	0.247	0.646
Norway	0	0.582	0.341	0.738	0.320	0.161	0.571
Spain	1	0.752	0.595	0.886	0.541	0.321	1.300
Sweden	0	0.709	0.534	0.844	0.505	0.276	1.025
Switzerland	3	0.510	0.289	0.681	0.257	0.176	0.422
UK	0	0.593	0.361	0.754	0.332	0.170	0.614

Notes: The number of lags is obtained using the t-sig information criterion of Ng and Perron (1995). $\hat{\alpha}_{MU}$ denotes the median-unbiased estimate of Andrews and Chen (1994). HL indicates half-life which is computed as $HL = \ln(0.5) / \ln(\hat{\alpha}_{MU})$.

Figure 1. Real exchange rates and the estimated structural breaks

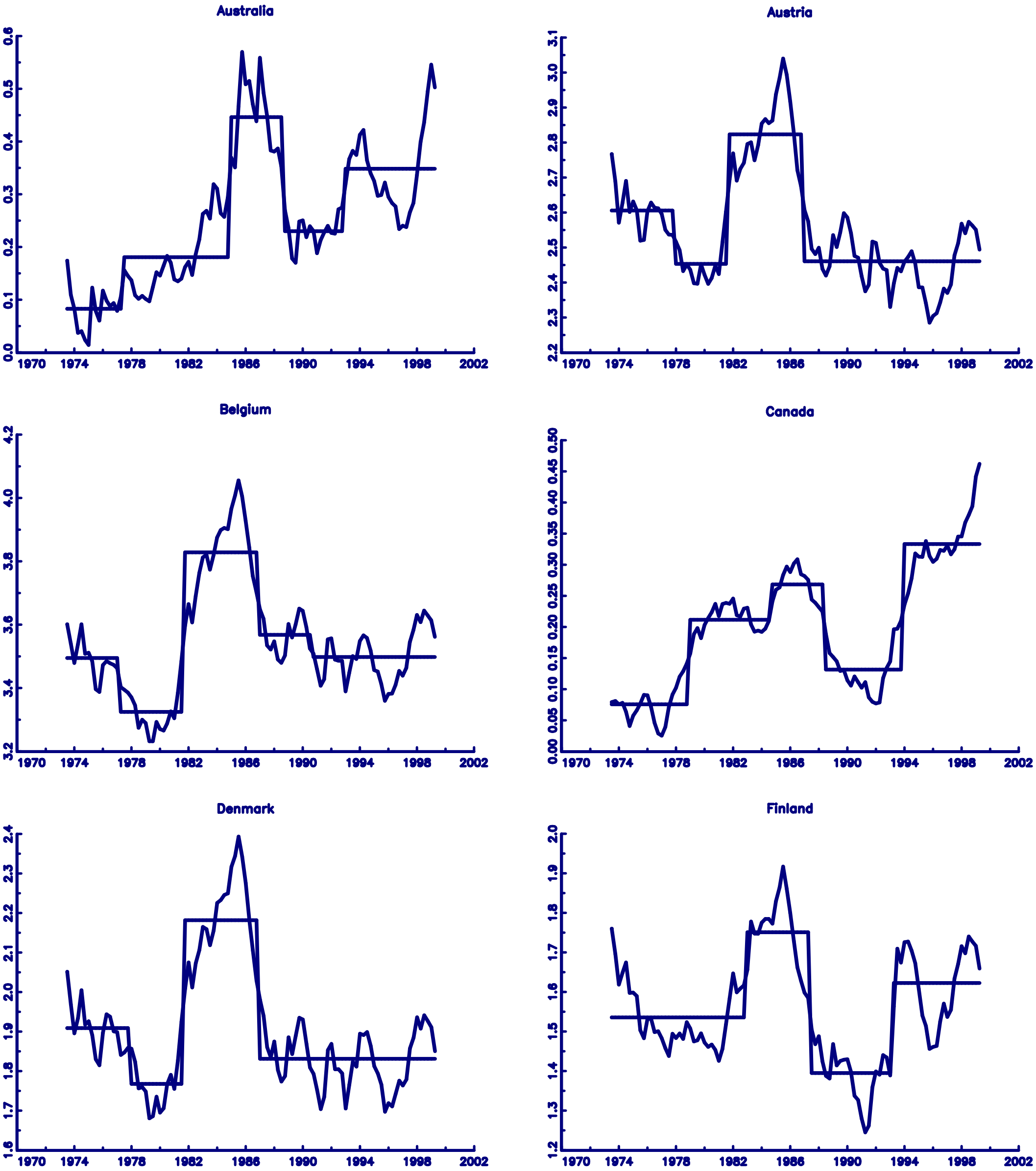


Figure 1 (Cont). Real exchange rates and the estimated structural breaks

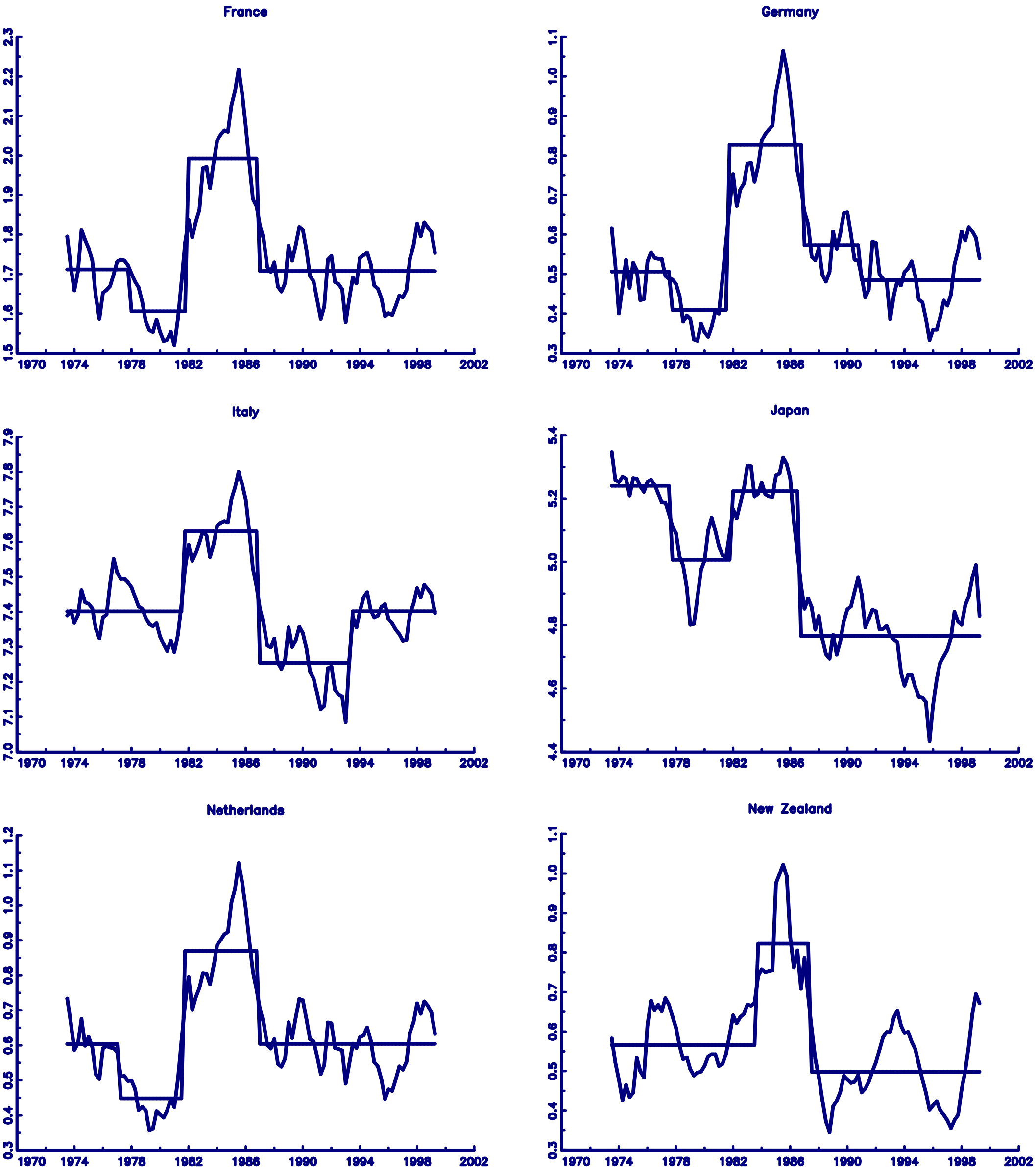
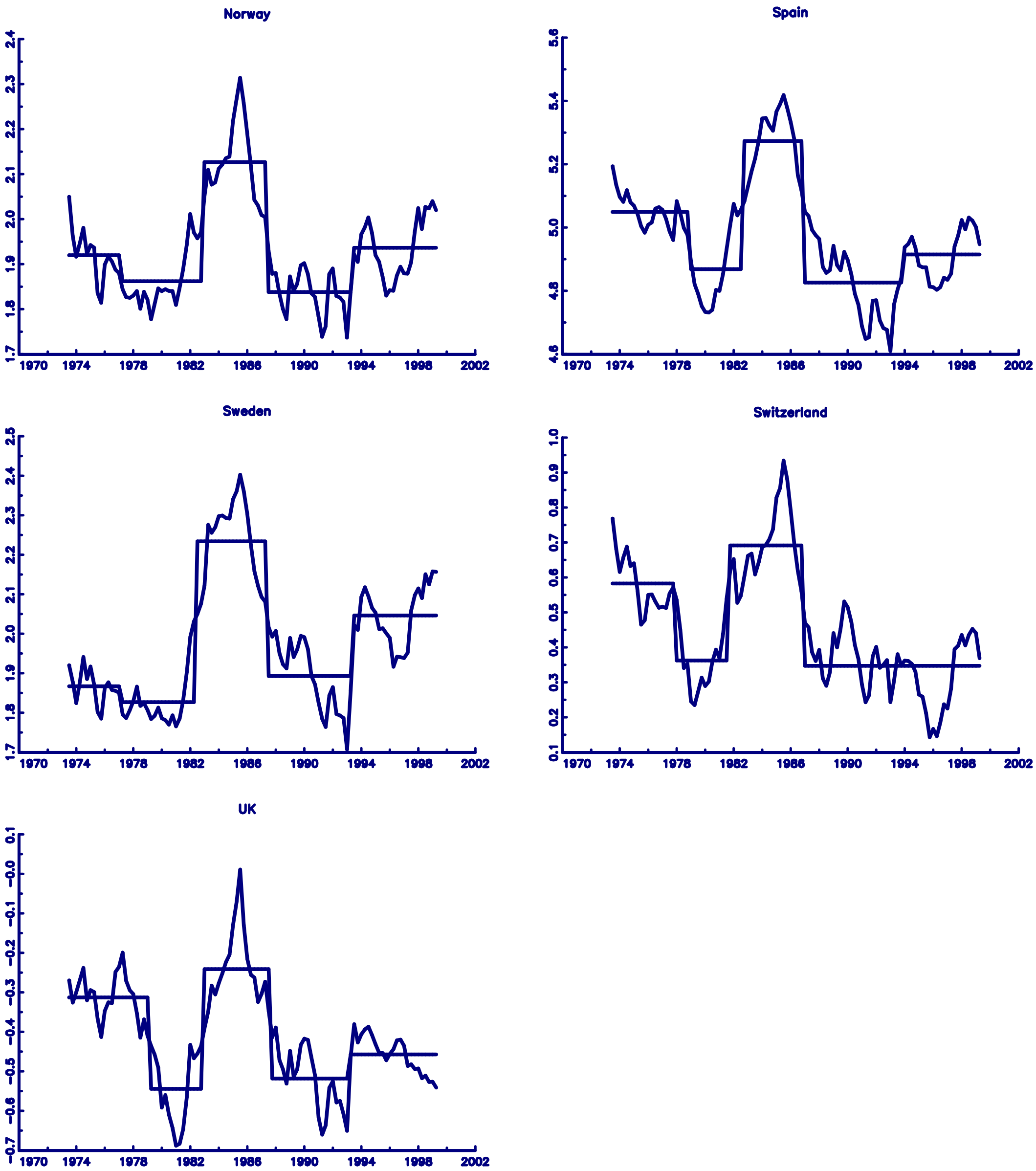


Figure 1 (Cont). Real exchange rates and the estimated structural breaks



2006

CREAP2006-01
Matas, A. (GEAP)**; Raymond, J.Ll.** (GEAP)
"Economic development and changes in car ownership patterns"
(Juny 2006)

CREAP2006-02
Trillas, F. (IEB)**; Montolio, D.** (IEB)**; Duch, N.** (IEB)
"Productive efficiency and regulatory reform: The case of Vehicle Inspection Services"
(Setembre 2006)

CREAP2006-03
Bel, G. (PPRE-IREA)**; Fageda, X.** (PPRE-IREA)
"Factors explaining local privatization: A meta-regression analysis"
(Octubre 2006)

CREAP2006-04
Fernàndez-Villadangos, L. (PPRE-IREA)
"Are two-part tariffs efficient when consumers plan ahead?: An empirical study"
(Octubre 2006)

CREAP2006-05
Artís, M. (AQR-IREA)**; Ramos, R.** (AQR-IREA)**; Suriñach, J.** (AQR-IREA)
"Job losses, outsourcing and relocation: Empirical evidence using microdata"
(Octubre 2006)

CREAP2006-06
Alcañiz, M. (RISC-IREA)**; Costa, A.; Guillén, M.** (RISC-IREA)**; Luna, C.; Rovira, C.**
"Calculation of the variance in surveys of the economic climate"
(Novembre 2006)

CREAP2006-07
Albalate, D. (PPRE-IREA)
"Lowering blood alcohol content levels to save lives: The European Experience"
(Desembre 2006)

CREAP2006-08
Garrido, A. (IEB)**; Arqué, P.** (IEB)
"The choice of banking firm: Are the interest rate a significant criteria?"
(Desembre 2006)

CREAP2006-09
Segarra, A. (GRIT)**; Teruel-Carrizosa, M.** (GRIT)
"Productivity growth and competition in spanish manufacturing firms:
What has happened in recent years?"
(Desembre 2006)

CREAP2006-10
Andonova, V.; Díaz-Serrano, Luis. (CREB)
"Political institutions and the development of telecommunications"
(Desembre 2006)

CREAP2006-11
Raymond, J.L.(GEAP)**; Roig, J.L..** (GEAP)
"Capital humano: un análisis comparativo Catalunya-España"
(Desembre 2006)

CREAP2006-12
Rodríguez, M.(CREB)**; Stoyanova, A.** (CREB)
"Changes in the demand for private medical insurance following a shift in tax incentives"
(Desembre 2006)

CREAP2006-13
Royuela, V. (AQR-IREA)**; Lambiri, D.**; **Biagi**, **B.**
"Economía urbana y calidad de vida. Una revisión del estado del conocimiento en España"
(Desembre 2006)

CREAP2006-14
Camarero, M.; Carrion-i-Silvestre, J.LL. (AQR-IREA)**.**;**Tamarit**, **C.**
"New evidence of the real interest rate parity for OECD countries using panel unit root tests with breaks"
(Desembre 2006)

CREAP2006-15
Karanassou, M.; Sala, H. (GEAP).;**Snower** , **D. J.**
"The macroeconomics of the labor market: Three fundamental views"
(Desembre 2006)

2007

XREAP2007-01
Castany, L (AQR-IREA)**; López-Bazo, E.** (AQR-IREA)**.**;**Moreno** , **R.** (AQR-IREA)
"Decomposing differences in total factor productivity across firm size"
(Març 2007)

XREAP2007-02
Raymond, J. Ll. (GEAP)**; Roig, J. Ll.** (GEAP)
"Una propuesta de evaluación de las externalidades de capital humano en la empresa"
(Abril 2007)

XREAP2007-03
Durán, J. M. (IEB)**; Esteller, A.** (IEB)
"An empirical analysis of wealth taxation: Equity vs. Tax compliance"
(Juny 2007)

XREAP2007-04
Matas, A. (GEAP)**; Raymond, J.Ll.** (GEAP)
"Cross-section data, disequilibrium situations and estimated coefficients: evidence from car ownership demand"
(Juny 2007)

XREAP2007-05
Jofre-Montseny, J. (IEB)**; Solé-Ollé, A.** (IEB)
"Tax differentials and agglomeration economies in intraregional firm location"
(Juny 2007)

XREAP2007-06
Álvarez-Albelo, C. (CREB)**; Hernández-Martín, R.**
"Explaining high economic growth in small tourism countries with a dynamic general equilibrium model"
(Juliol 2007)

XREAP2007-07
Duch, N. (IEB)**; Montolio, D.** (IEB); **Mediavilla, M.**
"Evaluating the impact of public subsidies on a firm's performance: a quasi-experimental approach"
(Juliol 2007)

XREAP2007-08
Segarra-Blasco, A. (GRIT)
"Innovation sources and productivity: a quantile regression analysis"
(Octubre 2007)

XREAP2007-09
Albalate, D. (PPRE-IREA)
"Shifting death to their Alternatives: The case of Toll Motorways"
(Octubre 2007)

XREAP2007-10
Segarra-Blasco, A. (GRIT); **Garcia-Quevedo, J.** (IEB); **Teruel-Carrizosa, M.** (GRIT)
"Barriers to innovation and public policy in catalonia"
(Novembre 2007)

XREAP2007-11
Bel, G. (PPRE-IREA); **Foote, J.**
"Comparison of recent toll road concession transactions in the United States and France"
(Novembre 2007)

XREAP2007-12
Segarra-Blasco, A. (GRIT);
"Innovation, R&D spillovers and productivity: the role of knowledge-intensive services"
(Novembre 2007)

XREAP2007-13
Bermúdez Morata, Ll. (RFA-IREA); **Guillén Estany, M.** (RFA-IREA), **Solé Auró, A**. (RFA-IREA)
"Impacto de la inmigración sobre la esperanza de vida en salud y en discapacidad de la población española"
(Novembre 2007)

XREAP2007-14
Calaeys, P. (AQR-IREA); **Ramos, R.** (AQR-IREA), **Suriñach, J**. (AQR-IREA)
"Fiscal sustainability across government tiers"
(Desembre 2007)

XREAP2007-15
Sánchez Hugalbe, A. (IEB)
"Influencia de la inmigración en la elección escolar"
(Desembre 2007)

2008

XREAP2008-01
Durán Weitkamp, C. (GRIT)**; Martín Bofarull, M.** (GRIT) ; **Pablo Martí, F.**
"Economic effects of road accessibility in the Pyrenees: User perspective"
(Gener 2008)

XREAP2008-02
Díaz-Serrano, L.; Stoyanova, A. P. (CREB)
"The Causal Relationship between Individual's Choice Behavior and Self-Reported Satisfaction: the Case of Residential Mobility in the EU"
(Març 2008)

XREAP2008-03
Matas, A. (GEAP)**; Raymond, J. L.** (GEAP)**; Roig, J. L.** (GEAP)
"Car ownership and access to jobs in Spain"
(Abril 2008)

XREAP2008-04
Bel, G. (PPRE-IREA) **; Fageda, X.** (PPRE-IREA)
"Privatization and competition in the delivery of local services: An empirical examination of the dual market hypothesis"
(Abril 2008)

XREAP2008-05
Matas, A. (GEAP); **Raymond, J. L.** (GEAP); **Roig, J. L. (GEAP)**
"**Job accessibility and employment probability**"
(Maig 2008)

XREAP2008-06
Basher, S. A.; Carrión, J. Ll. (AQR-IREA)
Deconstructing Shocks and Persistence in OECD Real Exchange Rates
(Juny 2008)

xreap@pcb.ub.es