The role of the yield curve for European monetary policy: some evidence pooling national pre-EMU data

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Abstract

This paper tries to ascertain whether the term structure of interest rates is a useful instrument for the European Central Bank (ECB). To this end, we employ individual country data for the Euro area. Using pooled and panel cointegration techniques we conclude that there is an equilibrium relationship linking the long and the short-run interest rates for both the individual countries and the panel as a whole. This can give support to the combined use of country and area-wide information in EMU monetary policy.

Classification J.E.L.: C32, F36.

Key words: monetary policy, European Monetary Union, expectations hypothesis, cointegration, term structure of interest rates, spread, panel.

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1 Introduction

In this paper we analyze the relationship between inflation developments and the short and long run interest rates in the Euro area. For this purpose, we use individual country information for the pre-EMU period. The paper contributes to two strands of the literature. First, it is connected to the leading-indicator literature of the term structure (Mishkin, 1990b, 1990a, 1991) that claims that the term structure of interest rates at longer horizons contains useful information about inflation changes. Second, it is also connected to the role of the term structure as an indicator of the monetary policy stance (Bernanke and Blinder, 1992, Blinder, 1998). Short-term rates are directly influenced by monetary policy authorities, whereas long-term rates are generally market-driven and their reaction to everyday policy actions is sluggish.

From 1999 the Eurosystem is taking responsibility for monetary policy in the Euro area. The primary policy objective, price stability, was quantitatively defined and a wide range of economic indicators was set for the assessment of future price developments and the detection of risks to price stability. The Eurosystem has announced that it relies heavily on two *pillars* in determining the appropriate level of interest rates. The first of them is a money-growth indicator. The second pillar is essentially an inflation forecast. The prominence given to the money-growth indicator has been criticized by many observers, like Gerlach and Svensson (2003), claiming that the inflation forecast is likely to be more useful in assessing risks to price stability. Within the range of indicators considered by the Eurosystem as possible drivers of inflation, special attention deserves the use of the yield curve, that has been traditionally considered a channel for the transmission of monetary impulses to the real variables.

The role of the term structure, either as an intermediate target or as indicator of policy stance has been already analyzed in the European case using mainly aggregate variables[1]. Berk and Bergeijk (2000), for example, discuss the role of the yield curve as an information variable for the Eurosystem, whereas Estrella and Mishkin

(1997) conclude that the spread is a useful piece of information for inflation and output forecasting in Europe and that the term structure has a role in the European monetary policy. However, empirical evidence using individual country data for the Euro area is more scarce. Angeloni et al. (2002) suggest that the area-wide evidence should be complemented with an assessment build up from the country level. This point has been recently stressed by De Grauwe and Sénégas (2003). These authors identify two main problems associated with the implementation of a common monetary policy in the Euro area. The first one is the asymmetry in the transmission of monetary policy measures by the ECB[2]. The second problem is the uncertainty about the transmission process[3]. In order to deal with the above mentioned problems and to implement an optimal European monetary policy, it is not sufficient to use area-wide (Euro) data. They claim that it may be better to complement aggregate information by also using non-aggregated national data from the member countries.

In addition, the quality of the forecasts used in the implementation of monetary policy can be affected by the nature (national or union-wide based) of the variables. Marcellino et al. (2003) compare the predictive power of two forecasting methods for EMU variables, either by *pooling country-specific forecasts* or by *directly forecasting the aggregate variables using other aggregate variables.* In general, the pooling method provides better results in terms of accuracy than the union-wide forecasts, making a case for the use of the national aggregate strategy (what they call NA) instead of the European aggregates (EA).

In this paper we use country data for the Euro area to gain insight on whether short-run interest rates affect both the long-run interest rates and the inflation rate through the term structure. This will give us an indication of the feasibility of using effectively the term structure either as a transmission channel of monetary policy or as an information variable of monetary conditions. Additionally, we provide some new evidence to the debate on the adequacy of using country-specific instead of area-wide information in the formulation of the Euro area monetary policy.

More specifically, we contribute to previous empirical literature in various respects: first, we specify and test a relationship linking long and short run interest rates and inflation rates using national pre-EMU country data in a panel including all the Euro area members[4]; second, using this specification we can analyze the long-run relationship pooling country-specific data (Pool Mean Group Estimators) and using individual country data (cointegration panel techniques); third, the econometric techniques allow us to test for nested competing specifications as well as for homogeneity restrictions on the long-run parameters.

The remainder of the paper is organized as follows. We first present, in section 2, the theoretical arguments explaining the possible role of the term structure in the implementation of monetary policy. In Section 3 we study the relationship between long and short run interest rates in the Euro-area countries for the period 1980:1-1998:4. Conclusions are reported in a final section.

2 The term structure as an indicator in monetary policy

The term structure reflects market expectations about future economic conditions. According to Estrella and Miskin (1997), the term structure spread is an indicator of the stance of monetary policy. In particular, a low spread reflects relatively restricted monetary policy because the spread is low when short-term interest rates are high relative to long term interest rates. At the same time, the term structure spread can play an important role as a leading indicator of real activity and inflation.

The most commonly accepted explanation for the link between interest rates with different maturities is given by the expectations hypothesis of the term structure. According to this hypothesis, long rates are mainly determined by expectations about future short term rates and, therefore, the slope of the term structure contains information about future short term interest rates.

As Kozicki and Tinsley (1998) point out, this characterization of the term structure lies on three assumptions. First, there exists a short-run interest rate, such as the EONIA in the Euro-area, which correctly reflects the monetary policy impulses coming from the central bank. This implies in turn that the short-term interest rate is under the central bank control. Second, according to the expectations hypothesis of the term structure, current and expected movements in the policy-controlled short-run interest rates are the main determinants of the term structure of bond rates. Finally, monetary policy affects the real economy since the long-run interest rates reflect the opportunity cost of investment and consumption.

Although modern monetary instruments tend to assure that monetary policy can readily influence short-term rates, long-term rates are generally market-driven and do not react hastily to everyday policy actions. Therefore the transmission mechanism from monetary policy actions to real economic activity will depend on the relation between short and long-term interest rates. This crucial link seems not at all as close as the expectations theory predicts.

The expectation hypothesis relates the yield on longer-term financial instruments to expected future yields on short term instruments. Following Campbell and Shiller (1987), in the case of pure discount bonds:

$$R_t^n = \frac{1}{k}\sum_{i=0}^{k-1} E_t R_{t+mi}^m + c_t^{n,m} \qquad n > m, \quad k = n/m \tag{1}$$

where R^n and R^m are the interest rates of a bond with a maturity of n and m periods respectively, $E_t R_{t+mi}^m$ is the expected value of the m periods bond yield i times m periods ahead and $c_t^{n,m}$ is a term premium that may vary with n, m, and t. Often the expectations hypothesis is expressed with $m = 1$. We can define $c_t^n \equiv c_t^1$ and the short term rate as $R_t \equiv R_t^1$. Then equation (1) states that the yield on an n-period bond is equal to a term premium plus the average of expected short rates up to $n-1$ periods in the future.

Kozicki (1998) argues that from this specification two theories justify the role of the term structure spreads to help predict future inflation. The first one justifies it because inflation responds to monetary policy actions and the term structure reflects the *stance of monetary policy.* In order to support this view, the expectations hypothesis can be rewritten with $m = 1$ as:

$$R_t^n = \frac{1}{n}\sum_{i=0}^{n-1} E_t R_{t+i} + c_t^n \tag{2}$$

Therefore, equation (2) equates long-term yields to an average of expected future short-term yields plus a risk premium. This means that market participants expect short-term rates to average in the future and for n sufficiently large this may smooth cyclical variations. According to this theory, as long-term rates reflect average short rates over a relatively long time interval, long-term interest rates may act as a benchmark for short-term yields comparison. For example, if monetary policy is relatively tight, short rates may be relatively high compared to the long-term rates, so that the spread is small or negative. In contrast, accommodative monetary policy may be reflected by short rates that are relatively low compared to long rates - a large term structure spread. In addition, although the spread falls when monetary policy is tightened, short-term yields are those that move closely with the rate that serves as monetary instrument. Long-term yields may react to policy, but rarely rise one-for-one with short-term rate increases. Thus, the term structure spread usually falls when monetary policy is tightened and therefore, the term structure spread helps predict inflation because it reflects the stance of monetary policy and economic variables respond to it. A low or negative spread predicts that in response to a tight monetary policy real activity will slow down and inflation will decrease. The opposite happens when monetary policy is accommodative[5].

A second theory on why the term structure spread may help predict inflation states that the spread reflects the *direction of future inflation changes*. From equation (2), the term structure spread (as the difference between the yield on a n-period bond and a one-period bond) can be decomposed à la Mishkin, in the sum of the expected real rate changes, the direction of the expected inflation changes and a term premium[6]. In the empirical literature, this approach has given rise to the test of the so-called *inflation-change equations*:

$$\pi_{t+1}^{k} - \pi_{t+1}^{m} = \alpha_{k,n,m} + \beta_{k,n,m}(R_t^n - R_t^m) + resid_{t+1} \tag{3}$$

where $\pi_{t+1}^{k} \equiv (1/k)\sum_{i=0}^{k-1} \pi_{t+1+i}$ is the k-period inflation rate from t to $t+k$, and $resid_{t+1}$ is the regression residual.

In order to provide a glance at the role of the interest rate spread in the Euro-area as an indicator of the monetary policy stance, Figure 1 shows the long and short-term interest rates for the period 1978-1998 in four selected Euro-area countries: Germany, the core of the European Monetary System (EMS); France and Italy, two countries that participated in the EMS since its creation (although only the former stayed after the 1992 crisis); finally, Finland, that joined the EU in 1995 and only then entered the EMS. The evolution of the spread in the four countries is different and driven not only by the monetary policy stance, but also by other factors and forces. The first remark that follows from the graph is that, from its entry in the EMS in 1979 until 1992, the short-term interest rate in Italy is always above the long-term rate with the exception of some months at the beginning of the eighties. The reason for this behavior is the anti-inflationary commitment of the Italian authorities derived from the EMS participation. During the same period, the German spread is positive, due to the accommodative character of monetary policy in Germany, also resulting from its anchor role in the EMS. The case of France is similar to the Italian, although less acute: the two interest rates evolve very closely (a tight monetary policy), with short-term interest rate peaks in some periods of time. The effects of a particular event, such as the German reunification, are noteworthy: the sharp increase in short-term interest rates in Germany is followed by France and Italy, even when the risk of recession may have required a more accommodative policy. In contrast, Finland (that was outside the EMS) suffered an asymmetric shock with the USSR decomposition. The large spread between long and short-term interest rates indicates the accommodative character of monetary policy in Finland at the beginning of the nineties. Meanwhile, monetary policy in France or Italy looked very tight when, in fact, short-term interest rates were raised to maintain the French Franc and the Italian Lira in the EMS.

All in all, a review of the empirical evidence on the expectations hypothesis will show immediately that it is far from being conclusive. For the United States Mankiw and Summers (1984), among others[7] rejected the expectations hypothesis. In contrast, Hall, Anderson and Granger (1992) got favorable evidence on the expectations hypothesis. For the European countries the empirical evidence is also mixed and somewhat country-specific[8].

Moreover, there is also a part of the literature that tries to assess the predictive power of the spread for future inflation. These are the cases of the seminal works by Fama (2001) and Mishkin (1990b, 1991)[9] and, for the European case, the studies by Berk and Bergeijk (2000) and Estrella and Mishkin (1997).

In the next section we analyze the role of the term structure in the individual Euro area countries as an indicator of the monetary policy stance. Using cointegration tests and estimation techniques applied to panel data, we assess the long-run relationships between short and long-term interest rates and inflation. Moreover, we explicitly test for cross country homogeneity in the long-run relationship linking short and long-term interest rates. This issue is of special relevance to assess the feasibility of using the yield curve as a leading indicator of inflation (as an alternative to other macroeconomic indicators) by the European Central Bank.

3 Cointegration analysis of the term structure of interest rates in the Euro area. Pooled and panel analysis

The existence of cointegration relationships between short and long-run interest rates and the usefulness of the spread as an indicator of monetary policy stance in the EMU are analyzed using two complementary approaches. We apply the Pooled Mean Group Estimators technique by Pesaran et al. (1999) and the recent cointegration tests derived for panels to study whether the spread holds for each individual country, comparing the results of homogeneous and heterogenous[10]panel estimations.

The data are quarterly observations covering the period 1980:1 to 1998:4. The source is the *International Financial Statistics* of the International Monetary Fund, and the variables are the call money rate (r_t), the ten-year bond rate (R_t) and the CPI-based inflation rate (π_t)[11].

The analysis is carried out using a panel approach that is compatible with the hypothesis of cointegration. We first test the specification of the panel using the Pooled Mean Group Estimators by Pesaran et al. (1999): first, whether inflation should be included in the long-run relationship linking long and short interest rates and, second, the hypothesis of homogeneous slope parameters for all the countries.

Then, both homogeneous and heterogeneous panels are estimated to test whether the interest rate spread is a valid stationary relationship for both individual countries and the area as a whole. The comparison of the slope coefficients will allow us to gain further insight on the interest rate channel of the transmission mechanisms of monetary policy.

In what follows, we first study the order of integration of the panel variables and then the long-run links between the two interest rates. Next, once we have found evidence in favor of cointegration, we assess the stationarity of the spread.

3.1 Order of integration of the variables

In this subsection, and previous to the analysis of long-run relationships, we present the results obtained from the analysis of the order of integration of the variables using panel unit root tests[12]. We have applied the LM test for the null of stationarity proposed by Hadri (2000) with heterogeneous and serially correlated errors.

We present the unit root test results for the null of stationarity in table 1. We use the two statistics proposed by Hadri (2000), that are the panel equivalents to the Kwiatkowski et al. (1992) statistics for the time series case: the statistic Z_μ tests for the null of level stationarity, whereas Z_τ tests the null of trend stationarity against nonstationary alternatives. The two statistics proposed by Hadri (2000) are distributed as $N(0,1)$[13].

In table 1 the statistics computed for $long_t$, $short_t$ and $inflation_t$ are very significant for both model specifications, so that the null hypothesis of stationarity can be easily rejected. Thus, the three variables in the panel are $I(1)$.

As a conclusion, the panel unit root tests support the non-stationarity of the variables analyzed.

3.2 Pooled Mean Group estimation of dynamic panels.

The Pooled Mean Group (PMG hereafter) estimator proposed by Pesaran et al. (1999) combines two procedures that are commonly used in panels and that make this technique specially suited to analyze the Euro-area. First, the Mean Group (MG) estimator: separate equations are estimated for each group and then the Mean Group estimator is computed giving consistent estimates of the average of the parameters. However, this estimator does not take account of the fact that some parameters may be the same across groups. Secondly, the traditional pooling estimators (such as the fixed and random effects estimators), that allow the intercepts to differ across groups whereas all the other coefficients and the variances are constrained to be the same. Thus, the PMG estimator involves both pooling and averaging. This estimator allows the intercepts, short-run coefficients and error variances to differ freely across groups, but the long-run coefficients are constrained to be the same. However, an interesting feature of this methodology is that some of the long-run parameters can be also unconstrained, so that they may be different for each group. This possibility can be tested using LR-type tests.

In this case, we estimate and compare, in the form of error correction models, two competing specifications: the first one includes the long and short interest rates together with the inflation rate; in the second specification, the inflation rate is excluded from the model[14]. The implied long-run relationships are the following:

$$\begin{aligned} Model\ 1 &: long_{it} = \alpha_i + \beta_{1t} short_{it} + \beta_{2t} inflation_{it} + \varepsilon_{it} \\ Model\ 2 &: long_{it} = \alpha_i + \beta_{1t} short_{it} + \varepsilon_{it} \end{aligned}$$

The inclusion of inflation is justified in order to analyze the short-run dynamics of the relationships and as a test for the inflation change equation as commonly appeared in the literature.

The results obtained from the estimation of the above specifications are presented in table 2. The lags of the variables (that are set equal to two) have been selected using the AIC criterion. In the case of the model including inflation, the null hypothesis of cross-country homogeneity of the short-run interest rate and inflation coefficients is rejected using a LR test. In a second step, when we just restrict the short-run interest rate and allow for heterogeneity in the inflation coefficients, the null cannot be rejected. However, even allowing for heterogeneity, the inflation rate was not significant in the long-run relationship resulting from the estimated panel, nor in the dynamics of the error correction model of the panel. Therefore, a inflation-change equation is not supported in the panel including ten euro members. However, it should be noted that significant parameters for the acceleration of inflation were found in the cases of Austria, Finland and Italy[15]. These three countries have in common that the exchange rate commitment in the Exchange Rate Mechanism (ERM henceforth) did not exist during a part of the sample. It can inferred from the analysis that the ERM may have blurred the functioning of the link between the spread and changes in inflation for the rest of the countries. Gerlach and Smets (1995) suggest that exchange rate pressures in regimes with pegged exchange rates may have obscured the information in yield spreads about future interest rate changes. The limited information content of the term structure has also been found in individual country-studies, such as Jondeau and Ricard (1997) in the case of France and Koedjik and Kool (1995) in Germany[16]. For the Euro area, Berk and Bergeijk (2000) also claim that the practical usefulness of the yield spread is rather limited.

In the second model we exclude the inflation rate from the specification. The first important result is that both the AIC and the SBC criteria recommend the second specification. In this case also, homogeneity in the short-term interest rate is accepted. The estimated coefficients are shown in table 2 under the heading *PMG estimation results* together with their Student's t. Both are very significant: the short-run interest rate estimate is 0.654, whereas the error correction parameter with a t-value of -4.98 passes the cointegration tests as described by Banerjee et al. (1998) and recently tabulated in Ericsson and MacKinnon (2002).

Once the restricted specification (where the inflation rate is excluded from the model) is the one chosen in the PMG estimation, we are going to analyze the hypothesis of cointegration in a panel setting.

3.3 Panel cointegration test results: homogeneous panel.

In this section we will first apply the panel cointegration tests and estimation procedures for homogeneous panels to the relationship linking long and short-run interest rates. In this framework, that means that we allow for fixed specific effects for each country but restrict the slope coefficients to be equal for all the members of the panel. Kao (1999) proposed DF-type panel non-cointegration tests based on the OLS residuals from the homogeneous panel regression.

The DF test from Kao (1999) follows the model:

$$y_{it} = \alpha_i + \beta x_{it} + e_{it}, \qquad i = 1, ..., N, \quad t = 1, ..., T \tag{4}$$

where both y_{it} and x_{it} are random walks. Thus, under the null hypothesis of no cointegration, the residual series e_{it} should be non-stationary. The limiting distributions are asymptotically normal at mean zero. Kao proposes four Dickey-Fuller (DF) tests[17], as well as the augmented version (ADF) of the test.

We present in table 3 the results of the different tests[18]. According to them, we can reject the null hypothesis of no cointegration with the five tests, as the statistics are normally distributed.

The parameters obtained from the bias-adjusted OLS and $DOLS$ estimation are also shown in table 3. In the two cases, the coefficient for the short-term interest rate is highly significant and of the correct sign. However, there are some differences in the magnitude of the parameters, being the one corresponding to the $DOLS$ procedure 0.75 versus 0.67 from the adjusted OLS. We should note that the PMG

estimate of the parameter was also very close to these values, with a coefficient value of 0.65. The fit is better in the Dynamic OLS estimation, as it can be seen in table 3.

3.4 Panel cointegration tests: heterogeneous panel.

In this section, the parameters are allowed to differ across the cross-sections, so that we will analyze the so-called heterogeneous panel. Two types of tests are presented, with different null hypotheses. First, we compute the ADF test proposed by Kao (1999) for both the individual members of the panel and the whole panel. The second is a LM test that has cointegration as the null hypothesis.

Kao (1999) ADF test for varying slopes and intercepts, is based on the following model:

$$y_{it} = \alpha_i + x_{it}^{'}\beta_i + e_{it}, \quad i = 1, ..., N, \quad t = 1, ..., T \tag{5}$$

Here, each cross-section is estimated individually and the pooling from the panel is done in the final step where the panel test statistic is based on the average of the individual cross-section statistics. Thus, each cross-section is allowed its individual cointegrating vector. The cross-sections are then assumed independent of each other although heteroskedasticity across the cross-sections is allowed. The null hypothesis, based on the DF test applied to the error term, is written as $H_0 : \rho_i = 0$ and the t-statistic for each i is called t_{iADF}. In addition, McCoskey and Kao (1998) propose a residual-based panel test of the null hypothesis of cointegration, also called *panel* LM test.

Table 4 shows the results of the panel cointegration tests for heterogeneous panels. The individual and panel LM and ADF tests results are based on the $DOLS$ estimates for heterogeneous panels with 2 leads and lags. According to the *individual*

LM tests, the null hypothesis of cointegration cannot be rejected for the majority of the countries (the exceptions being Finland and France at 10% levels). Moreover, the *LM panel test* (-0.50) does not allow us to reject the null of cointegration at 5% (the critical value being 1.6449).

The *ADF individual tests* for the null hypothesis of non-cointegration are presented in the second column of table 4 under the heading *LM and ADF cointegration tests*. In this case, the null is rejected for all the countries in the sample, in the majority of the cases at 1% level of significance. Concerning the *ADF panel test*, the null can be also rejected at 1% level, finding, according to this test, strong evidence of cointegration.

Therefore, once the existence of cointegration has been assessed, both for the individual countries and the panel, we concentrate on the parameter estimates. The *DOLS* parameter estimates for a model with two leads and two lags are shown in table 4 under the heading *DOLS conintegration tests*, together with the $t-values$ in parentheses. It should be emphasized that this estimation method corrects for endogeneity and autocorrelation using parametric methods[19]. In table 4, the significant coefficients appear in bold. From the results, it should be stressed, first, that both the intercept and the slope parameter are significant in all the equations. In addition, the magnitude of these coefficients differs only slightly among the countries in the sample: the constant terms are included in the interval (1.513, 2.607), whereas the largest value of the short-term interest rate parameter is 0.874 (in the case of Belgium) and the smallest is 0.761 (Ireland).

The cross-country similarity of the coefficients confirm the results obtained in the homogeneous panel case, where the *DOLS* parameter was 0.75, showing that the restriction of common slopes does not seem to be too binding. Moreover, this evidence is also compatible with the results obtained in subsection 3.2. above using Pool Mean Group estimators. As an additional formal test for homogeneity, the

Wald-test of the homogeneity restriction applied to the panel has been carried out. This test is distributed as a $\chi^2(10)$, and has a value of 16.89 in this case, so that the null can be accepted (with a critical value of 16.92)[20]

Accordingly, this supports the similarity of the slope of the term structure for each individual country to the one obtained for the whole area, which we identified as the restriction of common slope. As an additional test, we can apply the Hadri (2000) Z_μ test (level stationarity) to the variable $spread_t$, that is the difference between the long and the short run interest rates for every country in the panel. Thus, we are imposing that the two rates are cointegrated with a $(1, -1)$ vector. The result is presented in the lower row of table 1, where the null hypothesis of stationarity cannot be rejected at 5%.

4 Conclusions

In this paper we use country pre-EMU data for the Euro area to gain insight on whether short-run interest rates affect both the long-run interest rates and the inflation rate through the term structure. This will give us an indication of the feasibility of using effectively the term structure either as a transmission channel of monetary policy or as an information variable of monetary conditions. The econometric methodology is based on two estimation techniques applied to panels of data to analyze the existence of long-run cointegration relationships: the Pool Mean Group Estimators, and the homogeneous and heterogeneous Dynamic OLS (DOLS) panel cointegration tests and estimates.

We add some new evidence to the debate on the adequacy of using country-specific instead of area-wide information in the formulation of the Euro area monetary policy. We contribute to previous empirical literature in various respects: first, we specify and test a relationship linking long and short run interest rates and inflation rates using national pre-EMU country data in a panel including all the Euro area members; second, using this specification we can analyze the long-run relationship pooling country-specific data and using individual country data in a panel; third, the econometric techniques allow us to test for nested competing specifications as well as for homogeneity restrictions on the long-run parameters.

Several conclusions can be drawn from the empirical results. First, the term structure seems to be a valid relationship both for each country individually considered and for the system as a whole. Second, the slope of the term structure for each individual country is fairly similar across countries. The cross-country homogeneity of the long-run relationships between short and long-term interest rates is of special importance, as many other alternative macroeconomic indicators that could be used by the ECB to monitor monetary policy are unlikely to be as homogeneous. This evidence opens the possibility of using the term structure in the Euro area, even if the information content was rather limited during the EMS period. This lack of information was partly due to the exchange rate commitment under the ERM and the move to inflation targeting strategies by the monetary authorities in many European countries. This is no longer the case under EMU. Moreover, recent literature argues that the informational content of the term structure increases with the deregulation and integration of the financial markets, that may deepen again with EMU. In addition, due to the homogeneity found in the short-long term interest rates relationship, the fears raised about the use of aggregates by the ECB if not discarded need to be, at least, qualified. Therefore, the combined use of country and area-wide information may be a sound strategy during the first years of EMU.

Notes

[1]See Angeloni et al. (2002) for an overview of the recent empirical literature for the euro area, either based on aggregate data or on individual country data.

[2]The asymmetry in the transmission of monetary policy for the European case has been studied in De Grauwe (2000) and Gros and Hefeker (2002), while the empirical evidence in the case of the United States supporting this view appears in Meade and Sheets (2002) and Heinemann and Hüfner (2002).

[3]The issue of the importance of uncertainty in the transmission of the monetary policy after the creation of EMU has been stressed, among others, by Dornbusch, Favero and Giavazzi (1998), Mihov (2001) and ECB (2001).

[4]With the exception of Greece (due to data unavailability) and Luxembourg, whose data is included in those of Belgium. We are aware that, like in any other paper using pre-EMU variables to derive consequences for the monetary union, we must hope that the dynamic relationships between the data have remained broadly stable even after the introduction of the euro. However, this heroic assumption is an unavoidable problem due to the fact that there is little data for the monetary union period.

[5]Note that other measures of the stance of monetary policy should also help predict inflation. According to Bernake and Blinder (1992), short-term interest rates move closely with the interest rate that serves as the instrument of monetary policy and might provide a better measure of the policy stance.

[6]See Kozicki (1998) for a detailed algebraic decomposition.

[7]See Shiller (1979), Shiller, Campbell and Schoenholtz (1983) and Campbell and Shiller (1991).

[8]See Camarero and Tamarit (2002).

[9]In addition, recent works are Jorion and Mishkin (1991), Frankel and Lown (1994), Engsted (1995), Koedijk and Kool (1995), Gerlach (1997), Davis and Fagan (1997), Schich (1999), and Tkacz (2004).

[10]Different slope parameters across the members of the panel.

[11]The same definitions are adopted in Estrella and Mishkin (1997)

[12]The unit root analysis has been performed using programs written in GAUSS.

[13]It should be stressed that these tests are, according to Hadri (2000), very adequate for series highly dependent over time with large T (the time dimension) and moderate N (the number of cross-sections).

[14]Y. Shin has made available the Gauss code to compute the PMG tests and estimates.

[15]These results are not presented in the paper but are available upon request.

[16]However, the evidence for Germany is mixed (see Gerlach, 1997).

[17]Kao constructs statistics whose limiting distributions are $N(0,1)$ and do not dependent on the nuisance parameters, that are called DF_ρ^* and DF_t^*. Alternatively, he defines a bias-corrected serial correlation coefficient estimate and, consequently, the bias-corrected test statistics and calls them DF_ρ and DF_t. According to Baltagi and Kao (2000), the main difference between the two groups of tests is that whereas the DF_ρ and DF_t tests are based on the strong exogeneity of the regressors and errors, the DF_ρ^* and DF_t^* are more adequate for cointegration with endogenous relationships between regressors and errors.

[18]The program NPT 1.3. by Chiang and Kao (2002) has been used to compute both the homogeneous tests and estimates, whereas the codes to compute the heterogeneous tests and estimates have been kindly provided by McCoskey and Kao.

[19]According to McCoskey and Kao (1998), the dynamic OLS estimators have better asymptotic properties than the fully modified and OLS estimators.

[20]See (2000) for a description of hypothesis testing in panel data cointegration regressions.

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Table 1: Hadri (2000) stationary panel tests

Variables	Z_μ	Z_τ
$long_{it}$	8.92*	172.70*
$short_{it}$	8.36*	169.26*
$inflation_{it}$	5.01*	64.08*
$spread_{it}$	1.83	—

Note: The statistic Z_μ does not include a time trend, whereas Z_τ does, and are normally distributed. An asterisk denotes rejection of the null hypothesis of stationarity. The number of lags selected is $l = 8$.

Table 2: Pesaran, Shin and Smith (1999) PMG estimation.

Comparison of the specified models:

Model 1: $long_{it} = \alpha_i + \beta_{1t} short_{it} + \beta_{2t} inflation_{it} + \varepsilon_{it}$

Model 2: $long_{it} = \alpha_i + \beta_{1t} short_{it} + \varepsilon_{it}$

		N=10		**Variables**	
	AIC	*SBC*	*LR test*	$short_t$	$infl_t$
Model 1	-1411.81	-1440.10	$\chi^2(18) = 33.80$ [0.013]	$= \forall$	$= \forall$
			$\chi^2(9) = 13.47$ [0.14]**	$= \forall$	$\neq$
Model 2	-1421.05	-1447.02	$\chi^2(9) = 11.31$ [0.25]**	$= \forall$	···

PMG estimation results

Variables	(N=10)	$short_t$	0.654	ecm_{t-1}	-0.088
			(6.58)		(-4.98)

Note: p-values in brackets and t-Students in parentheses.

Table 3: Homogeneous panel cointegration tests.

Kao (1999) DF and ADF tests

	Test	p-value
DF_ρ	-16.81^{**}	0.000
DF_t	2.17^{**}	0.015
DF_ρ^*	-30.76^{**}	0.000
DF_t^*	-5.67^{**}	0.000
$ADF(1)$	-5.29^{**}	0.000

Adjusted OLS and DOLS estimates.

Variable	Adjusted OLS	DOLS (2,2)
$short_{it}$	0.6767	0.7512
	(33.08)	(34.54)
R^2	0.69	0.86
$\bar{R}^2$	0.69	0.70

(a) The two asterisks denote rejection of the null hypothesis of non-cointegration at 5%. The tests statistics are distributed as $N(0,1)$.
(b) For the OLS and DOLS estimates, t-statistic in parentheses. The DOLS estimate corresponds to a model with two leads and two lags. Dependent variable: $long_{it}$.

Table 4: Heterogeneous panel cointegration tests.

MODEL: $long_{it} = \alpha_i + \beta_{1i} short_{it}$

LM and ADF cointegration tests.

Countries	LM test	ADF test
Germany	0.04043	-5.22^{**}
Austria	0.03246	-6.495^{***}
Spain	0.04904	-6.34^{***}
Finland	0.30504*	-5.12^{**}
France	0.25776*	-5.05^{**}
Netherlands	0.03293	-5.58^{***}
Italy	0.08688	-6.27^{***}
Belgium	0.04395	-5.16^{**}
Portugal	0.05164	-5.42^{***}
Ireland	0.05164	-7.19^{***}
Panel Tests	-0.50	-13.12^{***}

DOLS cointegration estimates.

Countries	intercept	$short_{it}$
Germany	**2.525**	**0.769**
	(3.92)	(12.26)
Austria	**2.342**	**0.784**
	(3.73)	(12.83)
Spain	**2.43**	**0.774**
	(4.05)	(3.84)
Finland	**2.066**	**0.802**
	(2.51)	(10.02)
France	**2.236**	**0.780**
	(2.63)	(9.47)
Netherlands	**2.083**	**0.829**
	(3.31)	(13.15)
Italy	**2.221**	**0.816**
	(3.16)	(11.45)
Belgium	**1.513**	**0.874**
	(2.13)	(12.33)
Portugal	**1.999**	**0.834**
	(3.46)	(14.46)
Ireland	**2.607**	**0.761**
	(4.17)	(12.43)

Note:
(a) The lags orders of the ADF tests are 1, whereas the DOLS estimates are obtained form a model with two leads and two lags.
(b) The tests and the models have been estimated using COINT 2.0. in GAUSS 3.0.
(c) The critical values at 1% (***), 5% (**) and 10% (*) for the LM test are 0.5497, 0.3202 and 0.2335 respectively for the case of one regressor (Harris and Inder, 1994).
(d) The critical values at 1% (***), 5% (**) and 10% (*) for the ADF test are -5.3587, -4.7423 and -4.4625 respectively from Phillips and Ouliaris (1990).
(e) For the DOLS estimates t-Students are reported in parentheses. Significant coefficients in bold.

Figure 1: Long and short-term interest rates. Selected countries.