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Interest rate pass-through in the EMU – new evidence using the nonlinear ARDL framework

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Abstract

This investigation puts the interest rate pass-through mechanism from policy to deposit rates in the EMU under closer scrutiny. By using the newly developed nonlinear ARDL framework of Shin et al. (2011), we are able to estimate asymmetric long-run as well as short-run coefficients. Previous studies usually assume symmetric long-run behaviour which is too restrictive according to our results. Based on fully harmonized data for three EMU countries and the EMU as a whole we disentangle short-run and long-run asymmetries as well as heterogeneity in the interest rate pass-through mechanism across EMU. Our results point to considerable asymmetries especially with regard to the long-run pass-through of money market rate changes as well as some heterogeneity between countries.

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1. Introduction

The transmission of policy interest rates to retail rates, usually referred to as interest rate pass-through (IRPT) mechanism, is a key element of the monetary transmission process. A faster and more complete transfer of changes in policy rates to retail interest rates increases the effectiveness of monetary policy with regard to the real economy. Especially for the ECB, the investigation of the IRPT mechanism is of peculiar interest as it can reveal heterogeneity across the monetary union which should be considered when taking policy decisions. As the ECB is concerned about the average inflation rate across all EMU countries, the ECB should know to what extent her policy rate changes are transmitted to retail rates across the countries of the monetary union. If there are considerable differences in the IRPT mechanism, this would complicate the conduct of monetary policy and the achievement of stable inflation rates all across the union. Consequently, a cross-country study of the interest rate pass-through mechanism for EMU countries is advisable. Accordingly, the IRPT has been thoroughly investigated by various authors. A comprehensive and neat survey of the literature is provided by de Bondt (2002, 2005).

While early studies as Cottarelli and Kourelis (1994) did not rely on cointegration techniques, more recent papers increasingly did so. For example de Bondt (2005) used the single equation method of Engle and Granger (1987) and the systems approach of Johansen (1988, 1991) to investigate the IRPT in the euro area. Most recent investigations which have been provided for example by Sander and Kleimeier (2004, 2006) and Hofmann (2006) also considered asymmetries in the PT mechanism. However, these papers analysed asymmetries only with respect to short-run dynamics, for example by investigating whether the PT differs between rising and decreasing policy rates. This view is probably too restrictive as only short-run asymmetries are considered. As a novel contribution, this paper tackles this issue by adopting the nonlinear ARDL (NARDL) approach of Shin et al. (2011) to empirically test if asymmetries in the IRPT mechanism for EMU countries both in the long- and short-run exist. Neglecting long-run asymmetries might bias the coefficient estimates and thus could be an explanation of the incompleteness of the IRPT usually described in the literature (Kwapil and Scharler, 2010). Furthermore, when comparing the IRPT across EMU, most studies did not rely on fully comparable data. To the best of our knowledge, only Vajanne (2007) and von Borstel (2008) use completely harmonized data.[1] Hofmann (2006) hints at this drawback that there are no fully comparable time series of sufficient length for EMU countries. In fact, there are only some authors that use at least partly harmonized data.[2]

With respect to our point, we could not find any other investigation that uses the nonlinear ARDL framework to compare the IRPT in EMU countries.[3] Nevertheless, not allowing for asymmetries both in the short- and long-run might be too restrictive. For example, one might assume that not each change in the policy rate is transmitted to retail rates due to menu costs or simply uncertainty about future developments of the money market rate. If a drop in money market rates is reversed soon, a bank that immediately adjusts its retail rates, would exhibit quite volatile credit or deposit rates which might be unattractive for customers. Furthermore, the magnitude of the PT might be influenced by the degree of competition in the banking sector. In a less competitive banking sector PT rates might be generally lower. Additionally, a reasonable assumption is that banks try to gain or maintain market shares in deposit markets. In order to do this, the PT might differ between interest rates hikes and cuts. To gain market

[1] However, Vajanne's (2007) focus in on β- and σ-convergence rather than estimating PT coefficients. Von Borstel (2008) does not consider asymmetries in her investigation due to her short sample period.

[2] For example Kok Sørensen and Werner (2006), van Leuvensteijn et al. (2008) or ECB (2009) use the MFI interest rate statistics for data since January 2003. Interest rates prior to 2003 are counted back.

[3] Greenwood-Nimmo et al. (2011) investigate the relationship between monetary policy rates and long-term rates for Germany and the US during the Great Moderation with the NARDL approach.

shares banks would have to pass-through increases faster or to a larger extent than decreases. Unfortunately, with our simple approach we are not able to disentangle these effects from the possibility that interest rate cuts give rise to excess reserves.

The investigation proceeds by first presenting the methodology in the next section. Afterwards, the data are described and section 4 presents the results. The last section concludes.

2. Methodology

We employ the newly developed nonlinear ARDL framework of Shin et al. (2011) to investigate the IRPT mechanism in EMU. This approach is a generalisation of the ARDL bounds testing approach of Pesaran et al. (2001). It allows for estimating asymmetric long-run as well as short-run coefficients in a cointegration framework. Due to the fact that interest rates are classified as nonstationary, simple OLS regressions would be biased ("spurious regression"). Transforming the data to first differences would of course solve the problem of spurious regression but by doing this one would remove important long-run information from the data. Breaking down the sample in periods of interest rate increases and decreases probably does not really seem to be an option because this would shrink the sample period and leaves us with too few observations for reliable estimates.

As a starting point, consider a potential long-run relationship between the money market rate mm_t and the bank rate br_t of the following form:

$$br_t = \alpha_0 + \alpha_1 mm_t^+ + \alpha_2 mm_t^- + \varepsilon_t \tag{1}$$

Thereby, $\alpha_1$and α_2 represent the asymmetric long-run coefficients describing the extent to which a change in the money market rate is transmitted to the retail rate. The money market rate is decomposed into its positive and negative partial sum: $mm_t = mm_0 + mm_t^+ + mm_t^-$. Accordingly, we suppose that IRPT might differ between increases and decreases of the money market rate.[4] The partial sums are computed as follows:

$$mm_t^+ = \sum_{i=1}^{t} \Delta mm_i^+ = \sum_{i=1}^{t} \max(\Delta mm_i, 0) \tag{2a}$$

$$mm_t^- = \sum_{i=1}^{t} \Delta mm_i^- = \sum_{i=1}^{t} \min(\Delta mm_i, 0) \tag{2b}$$

In fact, for our sample, the mean change of the EONIA is -0.02 and the median is 0.00 which supports the choice of zero as the threshold value. Furthermore, there are 48 months in which the EONIA decreased and 42 months in which it increased (in 17 months the EONIA did not change). Thus, our sample is broadly balanced with respect to the directions of EONIA changes. Moreover, this number of observations should be enough to estimate reliable coefficients for both regimes. However, of course, due to the large interest rate cuts triggered by the financial crisis the sum of EONIA decreases is larger in absolute value than the sum of increases.

As shown in Shin et al. (2011) one can extend (1) to the following nonlinear ARDL model:

$$\Delta br_t = \beta_0 + \beta_1 br_{t-1} + \beta_2 mm_{t-1}^+ + \beta_3 mm_{t-1}^- + \sum_{j=1}^{m} \gamma_j \Delta br_{t-j} + \sum_{k=0}^{n} (\delta_k^+ \Delta mm_{t-k}^+ + \delta_k^- \Delta mm_{t-k}^- + \mu_t \tag{3}$$

(3) nests the linear ARDL model presented in Pesaran et al. (2001) for the case of $\beta_1 = \beta_2$ and $\delta_k^+ = \delta_k^-$ for all k. Thus, (3) is less restrictive than a linear model.

The test whether a cointegrating relationship exists, corresponds to the null hypothesis $\beta_1 = \beta_2 = \beta_3 = 0$ in Equation (3). For this test, as its distribution is non-standard, Pesaran et al. (2001) tabulate critical values. Furthermore, restrictions of long- or short-run symmetry

[4] However, any other threshold would be possible.

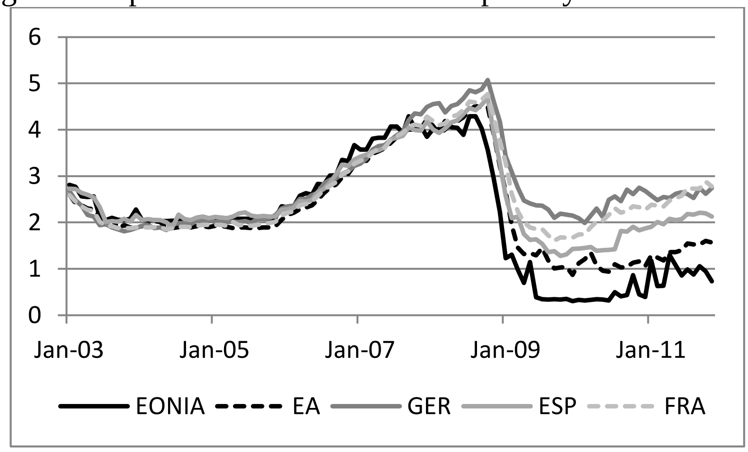
may be tested by conventional Wald tests. To obtain the long-run coefficients we normalize on the retail rate: $\alpha_0 = -\frac{\beta_0}{\beta_1}$, $\alpha_1 = -\frac{\beta_2}{\beta_1}$ and $\alpha_2 = -\frac{\beta_3}{\beta_1}$.

The lag length of the first differences of the retail rate and the money market rate in Equation (3) is chosen according to the Schwarz criterion where we consider a maximum lag length of twelve. In order to check whether the conventional ARDL bounds testing approach of Pesaran et al. (2001) is too restrictive, we estimate each model with this approach as well, applying the same methodology.

3. Data

We use deposit rates for households as well as enterprises stemming from the ECB's MFI interest rate statistics for Germany, France, Spain and the euro area as a whole. This choice was motivated by the fact that the time series were only completely available for these countries at the ECB's Statistical Data Warehouse Furthermore, an investigation for a couple of countries enables us to figure out whether possible nonlinearities are country-specific or a more general phenomenon. Moreover, according to the ECB Statistical Data Warehouse that publishes structural financial indicators, there is evidence that for example in France and Spain the banking sector is less competitive. It will be interesting to figure out whether these differences in competition will be mirrored in the coefficient estimates. Furthermore, this harmonized data set does not allow us to attribute different coefficient estimates to different data characteristics. To be precise, we investigate the PT for three different maturities – deposits up to one year, 1-2 years and over 2 years. With this, we can investigate if there are differences according to various maturities. All data are for new business. The sample runs from January 2003 until November 2011. Although this sample length might be at the lower limit when working with time series methods, we believe that using fully harmonized data to enhance comparability overcomes the caveat of only 107 observations per series. Additionally, this sample contains enough interest rate hikes and cuts to have enough observations for obtaining more or less balanced partial sums. The ECB's policy rate is approximated by the EONIA which was gathered via Datastream. We chose the EONIA as it reflects the stance of monetary policy most closely.[5] Additionally, using the same policy rate for each model improves the comparability of the results.

Figure 1: deposit rates for households up to 1 year and EONIA



[5] Thus, our approach could be labeled as "monetary policy approach". Another possibility would be to use the so called "cost-of-funds approach". This implies that the corresponding market rate is chosen according to the highest correlation with the retail rate under study. This emphasizes the funding cost of banks (Sander and Kleimeier, 2004).

In order to get an impression of the evolution of the series under investigation, Figure 1 displays the interest rates on deposits for up to one year for households. Furthermore, Table 1 displays some summary statistics for the deposit rates under study. From Figure 1, one gets the impression that deposit rates follow the EONIA quite closely. Thus, this is a first piece of evidence for a long-run, i.e. cointegrating, relationship. However, especially at the end of the sample there seems to be some cross-country differences as the German rates dropped less than the Spanish ones. Unfortunately, a thorough analysis of the effects of the recent financial turmoil on the PT mechanism is probably not really possible because of too few observations.

When comparing the summary statistics in Table 1 across countries, maturities and between households and enterprises it turns out that the time series do not differ that much between countries when one considers the mean. But there are some differences regarding the change between the beginning of the sample in January 2003 and the end in November 2011. While for the euro area as a whole and for France and Spain, deposit rates are in most cases lower at the end of the sample, in Germany rates increased in four out of six cases. One can take this as another indication for country-specific effects in the PT mechanism.

Table 1: descriptive statistics

	deposit rates for households			deposit rates for enterprises		
	≤ 1 year	1-2 years	≥ 2 years	≤ 1 year	1-2 years	≥2 years
	Mean					
EA	2.27	2.94	2.83	2.16	2.93	3.78
GER	2.76	2.79	2.54	2.55	2.87	2.74
ESP	2.51	3.15	2.91	2.30	2.85	3.17
FRA	2.61	2.90	2.72	2.30	2.95	3.43
	Variance					
EA	1.08	0.80	0.14	1.49	1.16	0.64
GER	0.81	0.71	0.20	0.94	0.80	0.33
ESP	0.80	0.32	0.13	1.36	0.97	0.47
FRA	0.77	0.64	0.15	1.22	0.85	0.43
	Maximum					
EA	4.51	5.00	3.87	4.47	6.08	5.29
GER	5.07	4.84	3.89	4.72	5.15	4.85
ESP	4.69	4.51	3.87	4.73	5.21	4.89
FRA	4.77	4.87	3.71	4.52	5.34	4.69
	Minimum					
EA	0.87	1.46	2.12	0.42	1.38	2.11
GER	1.81	1.72	1.80	1.25	1.42	1.46
ESP	1.28	2.09	2.45	0.56	1.32	1.99
FRA	1.61	2.01	2.04	0.66	1.85	2.22
	Change January 2003 – November 2011					
EA	-1.03	-0.89	-0.69	-1.67	-2.55	-2.15
GER	0.06	0.50	0.11	-0.52	0.54	-0.52
ESP	-0.60	0.58	-0.74	-1.40	-0.48	-0.20
FRA	0.19	0.29	-0.16	-1.26	-0.83	-1.46

None of the time series under study has to be classified as integrated of order two according to results of unit root tests.[6] Furthermore, the vast majority of time series is integrated of order one. However, both the ARDL and the NARDL approach are able to handle combinations of I(1) and I(0) variables (Pesaran et al., 2001). Of course, one might challenge the finding that interest rate are nonstationary from a theoretical point of view. Nevertheless, we will assume that they are for at least three reasons. First, the visual impression (see Figure 1) supports this notion as it does not seem that interest rates fluctuate around a constant mean for the period under investigation. Second, nearly all empirical applications in this field of research make the same assumption that interest rates are integrated of order one (see, for example, Sander and Kleimeier, 2004, 2006 or Hofmann, 2006). Accordingly, if we would make a different assumption about the time series properties, our results would not be directly comparable to other findings. Since one aim of our paper is to answer the question whether assuming long-run symmetry is too restrictive, this seems inadequate. And finally, we use a time series approach for estimation, i.e. we use unit root tests, to figure out whether the interest rate series

[6] For brevity, the results are not shown here but are available upon request.

are stationary or not. As all three points classify interest rates as I(1), we will follow this direction and use a cointegration approach to distinguish short-run from long-run effects.

4. Results

The detailed results of the NARDL approach as well as some diagnostics are shown in Tables 4 and 5 in the appendix, respectively.[7] To be able to derive some stylized facts about the IRPT, we test several hypotheses by conventional Wald test as indicated in the previous section.

H1: The IRPT mechanism can be described by a cointegrating relationship as shown in (1). This corresponds to rejecting the null hypothesis of $\beta_1 = \beta_2 = \beta_3 = 0$ in Equation (3).

H2: The IRPT mechanism can be characterized by long-run symmetry. Accordingly, we test whether $\beta_2 = \beta_3$ in Equation (3) holds.[8]

H3: The long-run PT is complete for a) positive and b) negative changes in the money market rate. This is investigated by testing a) $\beta_1 = -\beta_2$ and b) $\beta_1 = -\beta_3$ in Equation (3).

H4: The immediate PT of positive and negative changes of the money market rate is symmetric. This corresponds to $\delta_0^+ = \delta_0^-$ in Equation (3).[9]

To give a short overview of the results, Table 2 and 3 summarize the qualitative results of our hypotheses tests. Table 2 displays how often we are unable to reject the corresponding hypothesis.

Table 2: summary of hypotheses tests

	households	enterprises
H1: cointegration	11/12	11/12
H2: long-run symmetry	2/12	2/12
H3: long-run PT equal to 1	6/24	8/24
H4: immediate PT symmetry	9/12	10/12

To sum up the general findings, the IRPT mechanism in the countries under study is characterized by:

- a long-run relationship between the EONIA and deposit rates.
- no long-run symmetry. Thus, previous studies addressing this issues might use approaches that are too restrictive.
- a long-run PT from the EONIA to deposit rates that is one-to-one only in very few cases.
- short-run symmetry.

Table 3 presents a more thorough analysis of the long-run PT. As can be seen from Table 2, the PT coefficient is not significantly different from zero in only 14 out of 48 cases. However, this picture disguises one important finding. Actually, there are some cases in which the long-run PT is even larger than one indicating that changes in the money market rate transmit to deposit rates overproportionally. Nevertheless, for about half of the cases the PT remains incomplete.

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[7] We do not list the constant and all the lagged short-run coefficients to keep the presentation neat. Of course, all these coefficients are available upon request.

[8] Long-run symmetry implies that deposit rates react to EONIA increases and decrease in the same magnitude in the long-run. Accordingly, both long-run coefficients must be equal. For long-run symmetry it is not necessary that both coefficients are equal to one. For example, the relationship can be symmetric when there is an underproportional PT.

[9] There are indeed two ways to define short-run symmetry in Equation (3). The one way is the path we take here. The other would be to define short-run symmetry as equality of the sums of the short-run coefficients, i.e. $\sum_{k=0}^{n} \delta_k^+ = \sum_{k=0}^{n} \delta_k^-$. Of course, this test would imply a stronger definition of short-run symmetry. However, as our models contain different lag lengths as chosen according to the Schwarz criterion, we test short-run symmetry by the first approach.

Table 3: summary of long-run coefficients

	households	enterprises
Long-run PT >1	3	4
Long-run PT <1	15	12
Long-run PT =1	6	8

Taking now a more detailed look at the results it first turns out that when comparing the results of the conventional ARDL approach[10] and the NARDL framework, it is apparent that cointegration is found less often when inference is based on a linear ARDL model. When assuming linearity, cointegration is supported in only seven out of 24 cases. In contrast, when allowing for asymmetries, there are only two cases where we cannot reject the null hypothesis of no cointegration (see Table 2).[11] Moreover, the null hypothesis of no cointegration which corresponds to our *H1* hypothesis is soundly rejected in many cases. We take this as a first piece of evidence that neglecting nonlinearities might be too restrictive. Furthermore, with regard to the coefficient estimates, it is evident that the magnitude of the long-run PT is usually smaller in the linear case. Accordingly, the PT might be underestimated when not allowing for nonlinearities.

Turning to the coefficient estimates, we find that each long-run PT coefficient is considered to be significant. Furthermore, none carries the wrong sign and they all seem of reasonable magnitude. When comparing the magnitude between the positive and negative long-run coefficients a clear picture emerges. For all deposit rates but the rate for over 2 years for enterprises for the euro area the PT for the case of increasing interest rates is higher than for interest rate decreases. This finding could be attributed to competition in deposit markets. Banks pass-through money market rate increases to a greater extent to probably maintain or gain market shares. Or, to put it the other way round, banks pass-through EONIA decreases to a lesser extent not to lose clients. The Wald test of equality of both long-run coefficients is rejected in all but four cases. Thus, this difference in long-run PT is statistically significant and we have to reject our *H2* hypothesis.

Turning to *H3a)* and *b)*, which investigates the magnitude of the long-run PT in more detail, we find that both hypothesis are rejected for more than 50% of the models. However, the picture is different for both hypotheses. When *H3b)* is rejected, the PT is always significantly below unity pointing to an incomplete long-run PT of EONIA decreases. In contrast, the rejection of *H3a)* in most cases corresponds with a positive long-run coefficient of magnitude greater than one. Thus, increases of the EONIA are transmitted to the deposit rates overproportionally in the long-run. This again could be explained by the competition argument mentioned in the paragraph before and it supports the finding from *H2*.

With regard to *H4*, short-run symmetry is rejected in only a few cases. It seems that the immediate response to EONIA changes is passed-through in the same magnitude irrespective whether the EONIA rises or decreases. One reason for this finding could be that banks adjust their retail rates only if changes are large or if they believe that these are permanent. When this menu cost argument holds, the non-rejection of *H4* does not come as a surprise. The menu cost argument is supported by the finding of several insignificant short-run coefficients. Moreover, one should remember that we have tested only the weak version of short-run symmetry (see footnote 10). In the short-run structural determinants of the banking system might be less important. Comparing the magnitude of the short- and long-run PT, it becomes obvious that the short-run PT is generally lower than the long-run PT.

[10] The results of the linear ARDL approach are not shown here because the evidence of cointegration is so weak. However, the results are available upon request.

[11] Nevertheless, for completeness we present the results of these two cases here as well. Of course, we are aware that inference in both cases is not really correct.

Finally, we consider the PT across countries and maturities. When comparing the PT across maturities, the picture is quite clear cut: the PT decreases with maturity. The PT for deposit rates up to one year is in all but two cases the largest and the PT to deposit rates over two years is generally the smallest. This finding matches with the term structure of interest rates. Furthermore, for deposit rates with longer maturity the EONIA might not perfectly reflect funding costs. This could be further investigated by the “cost-of-funds approach” of the IRPT mechanism (Sander and Kleimeier, 2004).

The evidence of country specific findings is less obvious. The largest long-rung PT for deposits up to one year emerges for Germany. In Spain, the PT seems to be weakest. However, these findings do not hold for longer maturities. However, the finding that the PT seems to be highest in Germany can be attributed to structural financial indicators which are available at the ECB’s Statistical Data Warehouse. When considering the Herfindahl index which measures concentration in the banking sector (thereby, a smaller value corresponds to less concentration and more competition) it turns out that this index is considerably smaller for Germany compared to France and Spain. This notion is supported by the fact that the share of the five largest banks in Germany is only half as large on average compared to France and Spain. Accordingly, differences in PT rates can be explained by structural indicators.

Comparing the PT for households and enterprises, one result is that especially the negative long-run PT for deposits for enterprises is of larger magnitude. This impression holds for the positive long-run PT as well, although to a somewhat lesser extent.

5. Conclusion

The IRPT mechanism is of special interest for monetary policy makers as it allows drawing conclusions how fast and to what extent changes in policy rates are transmitted to retail rates. Therefore it has been investigated by numerous authors. However, possible long-run asymmetries have been neglected so far. This issue is addressed in this study. We consider long-run as well as short-run asymmetries in the IRPT in EMU by applying the newly developed NARDL framework of Shin et al. (2011). Our results strongly underscore that neglecting (long-run) asymmetries might bias the results. Not only that the evidence of cointegration is weak without allowing for an asymmetric long-run response of deposit rates to EONIA movements, moreover the long-run PT is generally stronger for EONIA increases than for decreases. Both findings could give rise to the assumption that competition on deposit markets works and banks try to maintain or even gain market shares by passing-through increases of markets rates to a larger extent than decreases. Additionally, the magnitude of PT declines with maturity and the positive PT for deposit rates up to one year seems to be even larger than unity. The evidence of heterogeneity with regard to deposits for enterprises versus households or between countries is less clear cut. Our results indicate that heterogeneity concerning EONIA increases and decreases is much more relevant. Nevertheless, we investigate the PT on a country level only for Germany, France and Spain. The picture could be less homogeneous when adding more countries of the EMU to our sample. However, the results for the deposit rates of the whole EMU do not differ that much from the single country results.

Our findings are quite relevant for policy makers. First, as the PT of policy rates differs between interest rate decreases and hikes, the change in the policy rate has to be stronger for expansionary policy measures than for contractionary ones to achieve the same stimulus for the real economy (in absolute values). Second, the magnitude of PT is probably larger than found in previous studies (Kwapil and Scharler, 2010). Finally, competition seems to work as policy rate hikes are transmitted more completely or even overproportionally to deposit rates. Furthermore, as the PT seems to be stronger in countries with a more competitive banking sector (i.e. Germany), policymakers should focus on ensuring a high degree of competition to keep the monetary transmission mechanism from money market to retail rates at work.

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Appendix

Table 4: results for deposit rates for households

deposit rate	β_2	β_3	δ_0^+	δ_0^-	*H1*	*H2*	*H3a)*	*H3b)*	*H4*	LM	ARCH
EA up to 1 y	1.169 (2.720)	1.002 (2.766)	0.392 (6.462)	0.072 (1.143)	5.461	0.020	0.009	0.952	0.002	0.000	0.468
GER up to 1 y	1.610 (1.957)	1.067 (1.883)	0.064 (0.771)	0.262 (3.652)	2.651	0.046	0.010	0.802	0.103	0.012	0.128
ESP up to 1 y	0.972 (6.426)	0.791 (6.288)	0.276 (3.576)	0.177 (2.138)	14.682	0.000	0.378	0.000	0.424	0.443	0.729
FRA up to 1 y	1.170 (4.010)	0.795 (3.800)	0.189 (3.636)	0.114 (2.233)	8.602	0.000	0.001	0.022	0.381	0.084	0.318
EA 1-2 y	0.913 (7.375)	0.777 (7.677)	0.217 (1.793)	-0.061 (-0.554)	20.272	0.000	0.150	0.000	0.135	0.406	0.649
GER 1-2 y	1.000 (5.578)	0.636 (5.537)	0.026 (0.159)	0.126 (0.875)	10.584	0.000	0.998	0.002	0.693	0.168	0.033
ESP 1-2 y	0.619 (4.288)	0.416 (4.217)	0.208 (1.420)	0.064 (0.519)	7.012	0.000	0.007	0.000	0.507	0.460	0.553
FRA 1-2 y	0.969 (7.347)	0.664 (7.315)	0.036 (0.381)	0.114 (1.321)	18.092	0.000	0.570	0.000	0.594	0.335	0.150
EA over 2 y	0.266 (3.318)	0.212 (3.628)	0.032 (0.208)	-0.243 (-1.832)	8.995	0.116	0.000	0.000	0.222	0.300	0.007
GER over 2 y	0.378 (4.918)	0.232 (4.755)	0.481 (2.600)	-0.236 (-1.512)	11.748	0.000	0.000	0.000	0.010	0.202	0.927
ESP over 2 y	0.339 (3.021)	0.260 (3.126)	0.054 (0.600)	-0.017 (-0.209)	5.771	0.121	0.023	0.007	0.601	0.049	0.071
FRA over 2 y	0.393 (5.713)	0.256 (5.595)	0.209 (2.151)	-0.180 (-2.043)	12.878	0.000	0.000	0.000	0.009	0.049	0.577

Values in brackets underneath the coefficients are t-values. The column labelled "*H1*" shows the F-statistic of the null of no cointegration (critical values according to Pesaran et al. (2001) are 4.78, 5.73 and 7.84 for the 10%, 5% and 1% level). For columns named *H2*, *H3a)*, *H3b)* and *H4*, see the beginning of section 4. The "LM" column gives the p-value of a test for autocorrelation up to lag 12. The last column contains the p-value of a test for ARCH effects up to lag 12.

Table 5: results for deposit rates for enterprises

deposit rate	β_2	β_3	δ_0^+	δ_0^-	H1	H2	H3 a)	H3 b)	H4	LM	ARCH
EA up to 1 y	1.084 (9.008)	1.017 (8.989)	0.313 (6.755)	0.316 (6.052)	27.343	0.000	0.000	0.087	0.973	0.100	0.105
GER up to 1 y	1.244 (2.411)	0.932 (2.285)	0.382 (4.791)	0.345 (4.608)	2.745	0.008	0.029	0.546	0.766	0.159	0.737
ESP up to 1 y	1.100 (8.692)	0.995 (8.632)	0.407 (5.184)	0.230 (3.008)	25.289	0.000	0.000	0.734	0.142	0.018	0.005
FRA up to 1 y	1.036 (3.750)	0.929 (3.761)	0.350 (7.541)	0.258 (5.286)	4.803	0.001	0.195	0.005	0.203	0.649	0.075
EA 1-2 y	1.125 (5.767)	0.929 (5.696)	-0.085 (-0.388)	0.527 (2.780)	11.213	0.000	0.128	0.277	0.064	0.243	0.245
GER 1-2 y	1.117 (4.760)	0.715 (4.645)	0.026 (0.116)	0.191 (2.790)	7.554	0.000	0.302	0.029	0.139	0.036	0.076
ESP 1-2 y	1.049 (6.196)	0.885 (6.413)	0.184 (1.279)	0.198 (1.630)	13.926	0.000	0.534	0.086	0.947	0.479	0.029
FRA 1-2 y	1.058 (5.698)	0.798 (5.758)	0.406 (2.335)	0.312 (2.067)	11.158	0.000	0.431	0.007	0.719	0.780	0.993
EA over 2 y	0.550 (4.619)	0.568 (5.096)	-0.361 (-1.291)	0.342 (1.451)	9.700	0.686	0.002	0.000	0.091	0.816	0.151
GER over 2 y	0.430 (4.709)	0.320 (4.827)	0.388 (1.232)	-0.304 (-1.154)	13.517	0.002	0.000	0.000	0.139	0.286	0.187
ESP over 2 y	0.775 (5.120)	0.601 (5.248)	0.372 (2.411)	0.091 (0.703)	9.510	0.001	0.054	0.001	0.213	0.672	0.717
FRA over 2 y	0.548 (5.022)	0.509 (5.203)	0.076 (0.392)	0.313 (1.903)	9.842	0.214	0.000	0.000	0.413	0.792	0.409

See Table 4 for explanations.