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The Determinants of Pass-Through of Market Conditions to Bank Retail Interest Rates in Belgium

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Editorial

On May 17-18, 2004 the National Bank of Belgium hosted a Conference on *"Efficiency and stability in an evolving financial system"*. Papers presented at this conference are made available to a broader audience in the NBB Working Paper Series (www.nbb.be).

Abstract

We analyse the pass-through of money market rates to retail interest rates at the disaggregate level in the Belgian banking market. First, we measure the extent of pass-through for a total of fourteen products. We find that the response varies over loans and deposits and depends positively on the maturity of the product. Second, the launch of EMU has generally not resulted in more competitive pricing by banks. Third, we assess the importance of several biases and find that heterogeneity in price-setting behaviour should be accounted for in analysing the pass-through. Fourth, we analyse bank-specific determinants of heterogeneous interest rate pass-through. We find a role for capital, liquidity and market share and we relate these results to the various channels in monetary policy transmission and to the structure-conduct-performance hypothesis in banking.

JEL-code : C23, E43, E52, G21, L11

Keywords: Pass-through, Heterogeneous panel, Aggregation bias, Panel cointegration, Retail-banking, Bank-level interest rates, Bank-level determinants

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1. Introduction

Recent years have been characterized by an increasing interest in the mechanics of the transmission of monetary policy to the banking sector. Even in the U.S., where bank finance is much less predominant than in the Euro area, the issue has attracted considerable attention (see e.g. Bernanke and Gertler (1995), Kashyap and Stein (2000)). The present paper elaborates on the issue by studying the pass-through of market interest rates to retail bank interest rates in Belgium, a member country of EMU. Cross-country evidence suggests that interest rate behaviour in Belgium is rather representative for EMU as a whole (see e.g. the overview in de Bondt (2002)). Our analysis focuses on 1) the measurement of the pass-through and 2) determinants of differences in pass-through among different types of banks. Both from a monetary policy and a banking perspective a thorough understanding of the pass-through is crucial. On the one hand, measurement of the pass-through provides insight into the extent and timing of agents' reaction to monetary policy, and market conditions more generally. Such analysis aids in the comprehension of, for instance, lags in the transmission of monetary policy. Investigating determinants of varying responses, on the other hand, can help in identifying the exact mechanisms at work in the banking sector. In this respect, pass-through research is complementary to both the analysis of interest spreads in the banking literature and the monetary transmission literature on the importance of the respective channels.

The present analysis contributes to the existing pass-through literature in a number of ways. A first contribution is related to the uniqueness of the data we use. Contrary to aggregate, country-level studies (e.g. Cottarelli and Kourelis (1994), Borio and Fritz (1995), Mojon (2000)), we have retail interest rates of individual banks at our disposal (see e.g. Cottarelli et al. (1995), Weth (2002), Gambacorta (2004)). This enables us to measure the extent of pass-through at the micro level and investigate its bank-specific determinants. We find that there is considerable heterogeneity in price-setting among banks. On the aggregate level, our results confirm the rigidity of prices, especially in the short run. Second, the analysis also comprises the liabilities side of retail banking. In contrast to the US (Hannan and Berger (1991), Neumark and Sharpe (1992)), there hardly exists evidence on the pass-through for bank deposits in EMU. Using data on Belgian banks, we account for some of this shortage. Third, we analyse a total of fourteen loan and deposit products, covering about the whole spectrum of retail banking activities. In contrast to existing studies, which typically consider only a limited number of products, this leaves scope for identifying pass-through characteristics over distinct product categories. Indeed, we find that corporate loans are priced more competitively than consumer loans. We also uncover a relationship between the product's maturity and the extent of pass-through that calls for a re-interpretation of some previous findings. Fourth, having bank balance sheet data allows investigation of factors driving differences in pass-through (Cottarelli et al. (1995), Bruggeman and Wouters (2001), Weth (2002), Gambacorta (2004)). We examine a more comprehensive set of determinants and find, among other things, that interest rates of well-capitalized banks are particularly sluggish in their adjustment to changing market conditions. Furthermore, our results also confirm the structure-conduct-performance, rather than the efficiency hypothesis.

Next to using microeconomic data, the paper also attempts to incorporate a number of methodological contributions in the setup of our empirical analysis. First, we argue that when dealing with retail bank interest rates, one should allow for heterogeneity. Not only fixed (as in Gambacorta (2004)), but also random variation over banks should be taken into account. Our results confirm the presence of both sources of heterogeneity. Second, we note that existing pass-through estimates potentially suffer from several biases. In particular, pass-through estimates might be contaminated by biases due to lagged dependent variables (e.g. Kiviet and Phillips (1994)), nonlinearity (e.g. Pesaran and Zhao (1999)), aggregation (e.g. Granger (1980)) or heterogeneity (e.g. Barker and Pesaran (1990)). Whereas previous studies often only mention these biases, we provide an indication of the importance of each of them. Our results indicate that heterogeneity in retail interest rate data is substantial, and that failing to account for this feature will give rise to misleading conclusions. On the other hand, we find no evidence to suspect that either lagged dependent variable bias or nonlinearity bias contaminates pass-through estimation. A final, but we hope very useful, contribution of our analysis lies in the way we test for and infer from cointegration relations. We incorporate lessons learned from the "large *n*, large *T*"-panel data literature in testing for (Pedroni (1999), McCoskey and Kao (1999)) and estimating (Phillips and Moon (1999)) cointegrating relations. On the one hand, we obtain a complete distribution of the long-run pass-through estimates. As a result, we shed light on the (in)completeness of the pass-through, an issue that has received little or no attention in the literature. On the other hand, applications of these techniques are not widespread, and almost exclusively cover macro data. In this respect, the paper also provides an illustrative (micro) example of some of the advances made in the "large *n*, large *T*" domain.

The remainder of the paper is structured as follows. Section 2 describes the structure of the dataset. In Section 3 we measure the pass-through. The specificity of the data has implications for the econometric approach used in the paper. After elaborating on the methodology, we present and discuss our results. Section 4 builds on the previous part and investigates which bank characteristics can account for heterogeneity in interest rate pass-through. A final section concludes.

2. The data

The dataset comprises bank-specific interest rates of the majority of Belgian banks for a series of loan and deposit products over the period January 1993 – December 2002. The sample period is interesting because it covers a period of six years preceding and four years following the start of EMU. From a macroeconomic perspective, the time span connects two episodes of weak, or even negative, economic growth, while the interval was characterized by relatively robust growth. In our sample period, the Belgian bank industry witnessed a pronounced shift in terms of market structure due to a series of mergers and acquisitions, which included all the major bank groups. Most large banks opted for the bancassurance model, causing the banking system to become more consolidated and more diversified. Both aspects may have an impact on the pricing behavior of retail bank products. Finally, the Belgian banking market is characterized by a relatively large degree of foreign penetration. Although the number of foreign banks is proportionately larger than their market share in various retail markets, it gives a rough indication of the market's contestability.

While aggregated bank retail rates are publicly available (see e.g. ECB, IFS), bank-specific rates are not. In Belgium, the central bank (National Bank of Belgium, NBB) conducts a monthly inquiry[1], asking banks what rate they offer or would offer on a range of fourteen standardized products[2]. The use of standardized products in the analysis has the advantage of eliminating to a large extent the effect of non-price competition. Moreover, since the inquiry covers high quality debtors, effects of time-varying default premia are also mitigated. Data are collected for six loan products and eight types of deposits, with both short and long maturities, and oriented both to consumers and firms. We refer to Table 1 for a description of the products and the abbreviations that are used for each of the products in the presentation of the results. The reporting banks account for more than 90% of (total assets of) Belgian banks[3]. For all but three[4] of the reporting banks we have the reported information, resulting in a total of 388 interest rate series. A second part of the dataset consists of money market interest rates of different maturities, which are publicly available. After determining the extent of pass-through from money market rates to bank-specific retail rates, we investigate the impact of bank characteristics on this relationship. We approximate bank characteristics by constructing ratios from banks' balance sheets and profit and loss accounts, which are also obtained from the NBB.

Prior to the analysis, some data issues need to be clarified. The first is the treatment of bank mergers or acquisitions that have occurred during the sample period. Some studies elect merging the interest rate data of merged banks into a single pre-merger series by retroactively computing an average interest rate over the whole sample. We do not follow this approach. First, such a strategy results in a loss of information. Second, prior to a takeover, it is unlikely that banks share the same pricing policy in their loan and deposit markets. Hence, since pricing policy is the object of our analysis, we treat banks that were involved in a merger or acquisition as different units before the merger, and as one thereafter. Regarding the bank retail rates, two adjustments are made. First, in January 1996, the definition for consumer credit in the NBB's inquiry changed, which we accommodate by restricting the analysis (for this product only) to the post-1995 period. Second, for mortgages, a (significant) dummy enters each regression to account for the level effect of an annual housing fair ("batibouw") affecting the interest rates reported in March. The frequency of the entire analysis is monthly. Finally, we restrict attention to those series for which we have at least thirty-six consecutive observations. The basic idea is that we require enough observations to estimate a time series model, although this requirement is somewhat less stringent when working in a panel context. A more technical reason is that having more time series observations than cross-sections allows usage of some particularly interesting inference procedures (see Section 3.b.2). The final dataset contains 31 banks and a total of 268 retail interest rates. The final dataset is unbalanced in the sense that different cross-section units within one panel are not required to cover the same time span. Furthermore, panels for different products may cover different banks.

We present our summary statistics in the form of a series of charts (see Figure 1)[5]. For each of the six lending products and the eight deposit products, we show a chart with the following

[1] While the inquiry is obligatory only since January 1996, the majority of banks reported voluntarily before that date.
[2] The products are standardized in the sense that maturity, amount and debtor quality are stipulated. The loan rates in the sample are typically those charged to the most creditworthy borrowers.
[3] Publicly available aggregate bank retail rates, e.g. available from the ECB, are constructed on the basis of these inquiries.
[4] From these banks we did not receive the required authorisation. It concerns three relatively small savings banks.
[5] In this case charts give more information than a table with simple summary statistics, because it would be difficult to capture both time variation and cross-sectional dispersion in a single table.

information: the evolution of the mean product-specific interest rate across all banks with reported rates in the sample, the evolution of the highest and the lowest product-specific rate charged by any bank in a given month and the evolution of the market interest rate with the same maturity as the specific product (dotted '+' line). Over most of the sample period from January 1993 to December 2002, interest rates have been declining, a characteristic that can largely be explained by the EMU-related convergence of interest rates and the stance of the business cycle and of monetary policy.

The general picture that emerges from the loan and deposit panels of Figure 1 is that bank retail rates generally follow changes in market rates, but there are clear differences across products in terms of speed and magnitude of the adjustment. On the lending side, we observe that the correlation between market and bank rates tends to increase with the maturity of the products (F5Y and C5Y are long-term corporate and customer loans, respectively). For short-term trade credit and bank advances (F2Ma and F2Mb), the interest rates offered by banks change only slowly and remain constant over various time intervals. For the long-term investment loans (maturity five years, F5Y) the association between market and bank rates appears to be more pronounced and the difference between the highest and lowest bank rate is also smaller. Next to the loan products offered to corporations, we also provide information on two typical customer retail lending products, mortgages (C5Y) and consumer credit (C3Y). For 36-month consumer credit, the bank rates appear to only loosely follow the corresponding market rate and the differences across banks are relatively high. In the case of mortgages the correlation between bank rates and the relevant market rate (5-year government bonds) is much more pronounced. However, there are considerable differences between the maximum and minimum bank rates, leaving scope to investigate bank-specific determinants of the pass-through. Especially in the first part of the sample period, the minimum mortgage rate tends to follow the benchmark government bond yield very closely, while the gap widens somewhat in the second part of the sample.

In the case of deposits, shown in the second panel of Figure 1, the association between market and retail rates is weakest for the savings accounts (SDA and SDL) and for demand deposits (DD). This can partly be explained by some typical features of the Belgian deposit market. It has been a longstanding practice of Belgian banks to offer a very low and stable interest rate on demand deposits, partly as a compensation for a series of payments services that were offered at low prices (cash cards, money transfer). In recent years competition among banks in terms of payment service fees has altered this picture. The case of savings accounts is even more typical because the interest rates are not subject to normal market forces. The maximum base rate on ordinary savings accounts is set by the Ministry of Finance. Customers that deposit their savings in these types of savings deposits are granted a partial exemption from the withholding tax. However, banks are free to offer higher base rates, but then customers have to pay the full withholding tax of 15%. Banks typically increase the potential interest rate on ordinary savings deposits by offering an accrual premium (for newly deposited money that is maintained on the savings deposit for a minimum time period) or a loyalty premium (on stable savings). From the corresponding charts, it is clear that the compensation for savings deposits is only broadly related to changes in market rates, often with considerable lags. In the case of savings bonds (SB1 and SB5) and time deposits (TD15D, TD3M and TD3Y), the evolution of bank retail rates is much more in line with changes in market rates. The differences between the maximum and minimum retail rates offered by banks are also much smaller than in the case of savings and demand deposits.

3. Measurement of the pass-through

In this section we analyse the extent of pass-through from market to retail bank interest rates; the next section deals with the determinants. We start by indicating the importance of heterogeneity in the analysis. This will result in an estimation strategy that will be used in this and subsequent sections. Next, we provide detailed information on model selection. We then present our results, compare them with the literature, and discuss some econometric issues such as potential biases and alternative estimators.

a) The estimation procedure

For each of the fourteen products, we consider a separate panel[6]. There are several characteristics of the data, of the hypotheses we wish to test, and of the underlying theoretical framework, that have implications for the way one should address estimation. We now turn to each of these.

As our study focuses on both the short and the long-term relation between bank retail rates and the market rate, we are interested in the time series characteristics of our data set. Having a cross-section of banks per product allows pooling these data and estimating the model more efficiently[7]. One should, however, be careful in pooling, as there may be considerable heterogeneity in the data. This heterogeneity seems especially relevant in our dataset and may be twofold: fixed (i.e. bank-specific) or random (i.e. uncorrelated with bank characteristics). First, differences in pass-through may be due to differences in bank characteristics. As this is one of the hypotheses we wish to test, allowing for fixed heterogeneity is crucial. We address issues concerning systematic heterogeneity in the next section, where determinants are dealt with. Second, recall that the interest rate data are the outcome of an inquiry, which allows for some subjectivity and/or differences in timing of reporting. Furthermore, the inquiry considers standardized products and not every bank may offer all the exact products. Allowing for random heterogeneity in estimation can capture the effect of measurement error in the data. Third, in regressions where bank characteristics are not included it may also mitigate the effect of fixed -but not modelled- heterogeneity. Fourth, another objection to pooling the data (especially in dynamic models like ours) is given by Pesaran and Smith (1995). They show that pooling in heterogeneous dynamic panels gives rise to inconsistent estimates. In Pesaran et al. (1999) an estimator is proposed that pools only in the long run, while allowing short-run heterogeneity. However, as there is no clear shared theoretical restriction for our cross-sections (banks), especially not when considering one product per panel, it is appropriate to allow for heterogeneity both in the short and the long run. Incorporating heterogeneity is feasible in our dataset as the time dimension is large (T ranges from 36 to 120). Hence, unlike the "large n, small T"-panel data literature, we are not forced to impose cross-sectional parameter equality. Moreover, the stationarity assumption present in that literature can be relaxed. Research concerned with "large n, large T"-panels has experienced a vast expansion since the mid-nineties. We discuss the implications of that literature for our modelling approach in greater detail in Section 3.b.

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[6] Examining cross-product relationships is beyond the scope of this paper.

[7] Throughout the paper, we refer to pooling or pooled coefficients whenever the estimation imposes common coefficients over cross-sections. Thus, in our terminology, the traditional pooled, fixed effects and random effects estimators all belong to the "pooled" class of estimators. The models we consider have fixed and/or random variation in all coefficients, not just in the constant term.

In practice, our estimation strategy has the following implications. When measuring the pass-through we estimate a pure random coefficient model. This implies that for bank i coefficients θ are of the form:

$\theta_i = \theta + \varepsilon_i$,

where $\theta = \Sigma\ w_i * \theta_i$. In words, individual coefficients are equal to the panel coefficient plus some random noise. Typically, the weights w_i are either equal, giving rise to the mean-group estimator proposed by Pesaran and Smith (1995), or a function of the respective group-specific estimated covariances, resulting in the Swamy estimator (see e.g. Swamy (1970), Hsiao (2003)).

b) Model selection and specification

1) Marginal costs

A first step in measuring the pass-through is determining the relevant marginal cost for each product. For almost all products in our dataset, the inquiry specifies a well-defined maturity. Hence, a natural choice for the marginal cost of the different products emerges: the money market rate with a comparable maturity. As Table 2 shows, this choice is largely confirmed by correlation analysis. We use the maximum correlation to determine marginal costs for those products for which a reference date was not specified. Table 1 presents all products and the chosen comparable money market rates. From an economic point of view, computing the pass-through relative to a market rate with comparable maturity rather than to the policy rate is important (see also de Bondt (2002)). Doing so, one disentangles the pass-through of marginal costs on the one hand, and term structure effects of policy rates[8] on the other.

2) Cointegration

The fact that a considerable part of our sample period is characterised by falling interest rates results in nonstationarity for the majority of series in our data set. For modelling purposes, in order to avoid spurious results, a natural question to ask is whether the respective retail and market rates are cointegrated. Overall, we follow the two-step procedure outlined by Engle and Granger (1987). Working in a panel context, however, alters the way one should test and estimate these models.

2.1) Testing for cointegration

It is common knowledge that standard unit root and cointegration tests have low power. The basic advantage of "large n, large T" panels[9] in using these tests is that exploiting the cross-sectional information improves their power. By now, the literature has established a wide variety of panel cointegration tests, surveyed in Banerjee (1999). Most of these tests differ in the specification of the null and alternative hypotheses. The arguments stated in Section 3.a indicate that we wish to allow for long-run heterogeneity. Therefore, we apply Pedroni's (1999) cointegration test, i.e. the between dimension ADF-test (see also Kao (1999) and McCoskey and Kao (1999)). This residual-based test under the null hypothesis of no cointegration allows for heterogeneity under the alternative. Moreover, the test does not

[8] The term structure effect of policy on market rates is typically not considered in pass-through studies, but is an interesting domain on its own. See e.g. Cook and Hahn (1989), Ellingsen and Söderström (2001).

[9] The cross-sectional dimension in our dataset is not large ($n \leq 31$) relative to the time dimension ($36 \leq T \leq 120$), but comparable to that of, for instance, McCoskey and Kao (1999). Furthermore, much of the theoretical results we rely on (e.g. Phillips and Moon (1999)) hold for moderate n and large T.

require the residuals of different cross-sections to have the same autocovariance process. Hence, interest rates are not required to exhibit a common behaviour over different banks.

2.2) Estimating and interpreting the cointegration relationship

When estimating a cointegration relation in the univariate case, OLS provides (super)consistent estimates. Due to the nonstationarity of the regressors, however, these estimators no longer have standard distributions. As a result, most of the pass-through literature avoids statistical inference on long-run coefficients[10]. One notable exception is Mizen and Hofmann (2004), who analyse long-run coefficients using likelihood ratio tests *vis-à-vis* the model with complete long-term pass-through. The present paper follows a different route, resulting in a complete distribution of the long-run coefficient.

In a panel context too, one can estimate the cointegration vector by OLS. A crucial difference with the univariate case, however, is that the noise in cointegration relations is attenuated when estimating over various cross-sections. As a result, Phillips and Moon (1999) show that panel estimators of cointegration coefficients converge to a normal distribution. Hence, when taking into account the appropriate distribution of the estimators, standard hypothesis testing becomes possible on long-run coefficients. The estimator Phillips and Moon (1999) suggest is a pooled estimator of the cointegration coefficient that can be interpreted as the average long-run coefficient of the heterogeneous individual cointegrating relations.

The outcome of the cointegration test determines the model that will be estimated. In case of stationary residuals, an error correction representation (ECM) of the retail rate exists:

$$\Delta b_{i,t} = c_{1i} + \sum_{k=1}^{p} \alpha_{ki} \Delta b_{i,t-k} + \sum_{l=0}^{q} \beta_{li} \Delta m_{t-l} + \gamma_i * u_{i,t-1} + \varepsilon_{i,t} \tag{1a}$$

where b = bank rate, m = market rate, $t = 1, \ldots, T$ and incorporation of heterogeneity is clear from the "i" subscripts on the parameters ($i = 1, \ldots, N$). The dynamic heterogeneity is captured by estimating *(1a)* using Swamy's procedure. The term γ_i* $u_{i,t-1}$ captures the adjustment towards equilibrium. A significantly negative γ_i confirms the presence of an equilibrium relation. The terms $u_{i,t-1}$ are the lagged residuals from individual cointegrating regressions, conform the Engle and Granger (1987) procedure:

$$b_{i,t} = c_{2i} + \delta_i m_t + u_{i,t}$$

In order to obtain one average long-run coefficient per product, the cointegrating vector is estimated directly using the pooled estimator suggested by Phillips and Moon (1999), which allows for long-run heterogeneity.

When a cointegration relation is absent, we estimate an autoregressive distributed lag model (ADL) in first differences, again using Swamy's random coefficient model:

$$\Delta b_{i,t} = c_{1i} + \sum_{k=1}^{p} \alpha_{ki} \Delta b_{i,t-k} + \sum_{l=0}^{q} \beta_{li} \Delta m_{t-l} + \varepsilon_{i,t} \tag{1b}$$

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[10] Remark that the distribution of the long-run coefficient affects the distribution of every intermediate, with the exception of the immediate, pass-through.

The final step in model selection is determining the optimal choice of lag length *(p, q)*. We use the Schwarz Bayesian Information Criterion to choose among models containing up to six lags (in levels)[11]. In the spirit of Granger causality tests, we first determine *p*, and then *q*. A moving average term is added to each equation to ensure white noise residuals and consistency of estimates.

c) Results

The results are summarized in Figure 2 and Tables 3, 4 and 5. Figure 2 plots the pass-through for the different products: the first panel contains the loan products, the second panel presents the deposit products. At each date, the pass-through measures the contribution of a 1%-point permanent increase in the market rate to the retail bank interest rate. Confidence intervals (95%) are computed from 5000 Monte Carlo draws[12], ruling out explosive roots (see e.g. Chen and Engel (2004)). Table 3 compares the random with the pooled coefficient model by means of likelihood ratio tests. The results always favour the random coefficient model. Table 4 shows the outcome of the cointegration tests, by means of the adjusted t-statistics for each panel. Table 5 summarizes the estimation results[13] and reads as follows. The table consists of two parts, one for loans and one for deposits. The first column identifies the specific loan or deposit product. Each coloured row contains the point estimates for one (product) panel. Within a panel, we restrict attention to the main coefficients of interest: the first three rows for each product report the immediate pass-through (ST PT), the long-run pass-through (LT PT) and the adjustment coefficient (ADJ). We also report the mean adjustment lag (mean lag) as a summary measure in the fourth row. Column 3 contains the actual point estimates and the adjacent column presents standard errors on these coefficients. In Column 5 we report how much of these standard errors is due to parameter heterogeneity. The remaining columns will be explained in the section on robustness checks.

First, for both loans and deposits, the size and significance of the negative adjustment coefficients (when estimated) confirm the presence of an equilibrium-restoring relationship. This is consistent with the results of the cointegration tests in Table 4. As the t-statistics in Table 4 show, the accrual premium on saving accounts is the only panel in which the null hypothesis of no cointegration is not rejected (t-statistic of 1.46). Although the cointegration test indicates the presence of a cointegrating relation in the demand deposit panel, its adjustment coefficient is only marginally significant[14]. Conversely, in the case of trade credit (F2Ma) cointegration is confirmed only marginally by the panel ADF test, but its adjustment coefficient is highly significant. All other loan and deposit products are cointegrated with

[11] We also considered sequential likelihood ratio tests as a selection procedure. For six of the fourteen panels, this resulted in the same lag choice. Six other panels had a different lag specification, but did not imply significant deviations from the Schwarz model. The remaining two panels had some of the intermediate responses lying outside the Schwarz model's confidence bounds, but never the main coefficients of interest. For the remainder of the paper, we consider the more parsimonious (Schwarz) model as our baseline specification.

[12] Although drawing from the distributions of the average parameters implies somewhat more computational effort, it is convenient in the sense that it does not require deriving standard errors for the pass-through -using e.g. the delta method- for each intermediate date. The resulting confidence bounds would not change very much.

[13] Starting the analysis from 1994 onward, excluding the late 1993 market turbulence due to the EMS-crisis, never had a significant impact on the results.

[14] This is also apparent in the lower confidence bound for this product: in the simulations, the adjustment coefficient often reaches a level close to zero, resulting in a very slow (or even no) adjustment towards the long-run relation.

their respective comparable market rates. Second, Figure 2 shows relatively large confidence intervals, for loans in particular. For the most part, this is due to the amount of cross-sectional variation in the pass-through estimates, as the fifth column of Table 5 indicates. We now discuss our results for the measurement of the pass-through in detail.

Let us consider the long-run pass-through for loans. Figure 2 and Table 5 show that the long-run response is one-for-one only for two of the corporate loans, *viz.* the term (F6M) and investment (F5Y) loans (their estimated long-term pass-through is 0.92 and 1.01, respectively, see Table 5). Recall that, in contrast to much of the existing literature, the completeness hypothesis is now truly tested. Regarding the point estimates, the long-run pass-through is relatively low for the two consumer-oriented products (69% for consumer credit and 87% for mortgages, products C3Y and C5Y respectively). The fact that the pass-through is often incomplete in the long run warrants caution in interpreting results stemming from estimations imposing complete long-term pass-through. Conform the bulk of evidence in the literature, we find considerable short-term stickiness (see ST PT in column 3 of Table 5), although there are differences across the loan products. Just as in the long run, this stickiness is most pronounced for the consumer loans in our sample, while only to a lesser extent for firm loans. Regarding consumer loans (C3Y and C5Y), at most 40% of the long-term pass-through is adjusted on impact, whereas for firm loans (F6M and F5Y) at least 80% of the final response is immediately realized. With respect to the speed of adjustment, computation of the mean lag reveals a similar result. Banks are slower in their adjustment to market rates for consumer-oriented products. An interesting result is that long-term adjustment seems more complete for both firm and consumer loans the longer their maturities. This result has not been identified in previous research due to either a lack of products to compare with, or the use of a short-term market rate, rather than one with comparable maturity. Pass-through estimates that do not distinguish between marginal cost and term structure effects -using the short-term market rate as marginal cost-, typically find the opposite: the pass-through is lower the longer the maturity of the product.

The observation that consumers are faced with less competitive pricing is consistent with the model of Rosen (2002). The latter argues that the more sophisticated (in terms of search intensity and access to alternative finance) customers are, the more complete the pass-through will be. Rosen (2002) finds evidence for his model using aggregate U.S. deposit data. Interpreting consumers as being less sophisticated than firms, we too find evidence in support of his model, but using disaggregated data on loans. The finding that the interest rate pass-through is faster and more complete for corporate loans than for consumer credit may have implications for the ability of the central bank to influence investment versus consumption decisions. In particular, the relatively weak pass-through of consumer loans may be a key element in understanding why the ECB seems to be incapable of influencing private consumption (Angeloni et al., 2003).

The other observation, *viz.* the longer the maturity of the product, the more complete the pass-through, is consistent with several theoretical considerations. First, from a general price-setting point of view, the presence of price staggering induces firms to also consider expected future marginal costs in setting today's prices. As these are not incorporated in the model, this might explain the finding of increased stickiness for shorter term products. For these products expected future marginal costs are, in line with the pure expectations theory, contained in long term interest rates. If the maturity effect would indeed be due to a misspecification of marginal costs, we should not find it when measuring the pass-through relative to a longer term money market rate. Thus, as a robustness check for the maturity effect, we also estimate the pass-through of the five-year money market rate to each product's interest rate. While point estimates (obviously) differ, the positive relation between the pass-through and the maturity of the product still stands[15]. Thus, we find that the maturity effect is robust to the specification of marginal costs. Secondly, the longer the maturity of a product is, the larger the scope for moral hazard phenomena to occur. In an attempt to avoid this, banks can follow the market more closely (reflected in a higher pass-through)[16]. Thirdly, products with longer maturities are typically those with large underlying amounts. The higher the amount of the loan, the more significant interest payments become for the customer. Hence, for these products the search for banks offering competitive conditions is more intense (Stigler (1961)). Increased customer search implies more competitive behaviour among banks, and thus a more complete pass-through.

Turning to deposits, an inspection of the estimates in the second panel of Table 5 reveals that we can distinguish two groups of deposit products with respect to interest rate pass-through: time deposits and savings bonds on the one hand, and current and savings accounts on the other. The second panel of Table 5 shows that complete long-run pass-through is always rejected statistically, although only marginally so for both the time deposit and savings bond with long maturity (the point estimates of 0.98 and 0.96, respectively, for products TD3Y and SB5 are close to 1). For time deposits and savings bonds point estimates of the long-run effect are found to be higher for longer maturities. Thus, liabilities seem to exhibit a similar maturity effect as the one found for loans. Although there is some immediate stickiness, adjustment to the long-run level is rapid. Table 5 shows that mean lags for time deposits and savings bonds are typically very low, mostly below 1.5 months. Demand and savings deposits, on the other hand, show a different picture. First note that the reward on savings deposits is threefold: it consists of a base rate augmented with a loyalty premium and/or an accrual premium. Figure 2 reveals that the accrual premium (plus the base rate) banks offer does not react in response to changes in money market rates. Marginal costs do play a role in the determination of the savings' loyalty premium. The response of the latter is very similar to that of demand deposits. Both these retail rates show considerable short-term stickiness. The estimated immediate pass-through is 6% for the loyalty premium and 14% for demand deposits. Even in the long run, the response is far from complete (63% and 49%, respectively, and these estimates are significantly different from one). Moreover, adjustment is particularly sluggish for both these products, as indicated by their mean lags (up to 7 months for SDL). Hence, the Belgian bank deposit market seems to consist of two distinct segments. The first is the market for time deposits and savings bonds, where banks seem to follow changes in market conditions quite rapidly in their competition for the deposits of companies and households. The second is the segment of current and savings accounts, where adjustments to market conditions are much slower and competitive pressure seems to be lower.

[15] For brevity, we do not report the exact results, but they can for the long-term pass-through largely be inferred from the correlations in Table 3. There it is apparent that considering expected marginal costs (in terms of longer term market rates) does not entail a more similar pass-through for long and short term products.

[16] The effects of asymmetric information have somewhat sharper predictions for asymmetry in pass-through.

d) Issues

This section and the accompanying appendices elaborate on the measurement results presented above. Firstly, as in de Bondt (2002), we analyse changes in pass-through after the introduction of the euro. Secondly, we describe the effect of using an alternative estimation strategy. Thirdly, all pass-through estimates potentially suffer from a bias due to either nonlinearity, the presence of a lagged dependent variable in the model, or both. We provide results based on some alternative estimators that correct for these biases. Finally, we investigate how our results relate to those in the literature that use different estimation procedures or data on a more aggregated level.

EMU results

After the completion of the EU single market in 1993, the introduction of the euro in 1999 was intended to be the final milestone in the creation of a truly integrated banking market in the Eurozone. However, studies on the remaining legal and cultural barriers (see Heinemann and Jopp (2002)) and empirical evidence on the pricing of retail banking products (see Corvoisier and Gropp (2002)) indicate that the degree of integration of Eurozone bank markets differs markedly between the wholesale, corporate and customer retail market segments. Within the context of interest rate pass-through, de Bondt (2002) and Sander and Kleimeier (2004) provide evidence suggesting less competitive behaviour after the launch of EMU. Similar to those studies, we re-estimate the pass-through for the EMU period (1999-2002) separately[17]. The results are presented in the right-hand panel of Table 5 and tend to confirm the weakening of the pass-through. Moreover, the variety of products in our analysis yields some additional insights.

Concerning the bank assets in our analysis, once again, a distinction should be made between firm and consumer-oriented products. Comparison of Columns 3 and 10 in Table 5 shows that the shift to EMU seems to have resulted in less competitive pricing of consumer products. From an economic point of view, the drop in both immediate and long-term pass-through is considerable for mortgages and consumer credit (e.g. the long-term pass-through for consumer credit is 0.58 in the EMU era compared to 0.69 for the full sample, which only covers three more years). We also observe an increase in the mean lag, altogether implying a slower adjustment to a lower long-run level. For firm loans, no such clear pattern is observed. Adjustments in long and short-term responses often go both ways. For the firm loans with very short maturity there seems to be an increase in sluggishness, while the others exhibit a minor drop in their mean adjustment lag. For liabilities, the results indicate a fall in long-term pass-through for all term deposits and savings bonds. This is combined with a rise in the short-term response for these products. Overall, the adjustment becomes quicker, but smaller. A remarkable response is found for the demand and savings deposits. In the EMU era, they no longer show any reaction to the respective market rates. The most common interpretation of these findings is that the Belgian corporate lending and deposit markets were already contestable and, hence, competitive in the pre-1999 period and that remains the case afterwards. On the other hand, the large banks, which dominate Belgian retail banking after the merger wave of the late 1990s, may use their market power to adjust their customer retail rates more slowly and less completely. Our pass-through results for the Belgian bank market lend no support to the conjecture that the single currency may have spurred more cross-border (actual or potential) competition.

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[17] This section contrasts the EMU subsample estimates with those of the whole sample. Comparing with the pre-EMU estimates does not change any of the conclusions.

Robustness: Alternative estimators
In Appendix A we show that the conclusions based on the measurement results presented earlier do not change when considering an alternative weighting of individual coefficients. Appendix B analyses to what extent our results suffer from a bias due to the presence of the lagged dependent variable in the model or due to nonlinearity of pass-through measures. Overall, the analysis and discussion of the respective biases presents some good news for both our analysis and possibly for pass-through research in general. Even though there may exist some minor differences in point estimation, broad conclusions about the measurement of the pass-through do not seem to suffer from lagged dependent variable or nonlinearity bias.

Aggregation bias
Since much of the empirical pass-through literature is based on aggregated data, it is of interest to know to what extent aggregate findings relate to those based on the micro data they stem from. Aggregation may bias estimation, especially in dynamic relations (Harvey (1981), Barker and Pesaran (1990)). To gain some insight in this respect, we also perform the analysis univariately on the aggregate Belgian ECB-retail rates over our sample period. Column 8 in Table 5 presents the results. Since the ECB data do not cover all fourteen products in our dataset, we can only provide estimates for the available comparable products. The good news for studies based on aggregate data is that differences in summary measures such as the mean lag seem limited. Turning to point estimates, differences become more substantial and less flattering for aggregate studies. Although some of the estimates are in line using either data (e.g. long-term pass-through for time deposits is 88% in both cases), lots of aggregate point estimates lie outside the confidence intervals surrounding the disaggregated estimates. We find systematic underestimation of adjustment coefficients when estimation is based on aggregate series. From an economic point of view, too, this is worrisome. Differences in (both long and short-term) pass-through often exceed 5%-points, reaching a maximum difference of 18%-points for the immediate pass-through of time deposits. Figure 3 graphically represents the systematic underestimation in adjustment coefficients. Thus, in contrast to the lagged dependent variable and nonlinear bias, aggregation bias may play an important role in pass-through estimations[18].

Heterogeneity bias
Appreciation of the possibility of aggregation bias has spurred pass-through research on disaggregate data. However, merely changing the level of aggregation of the data does not guarantee to solve the problem. The reason is that banks may well be heterogeneous agents. In this subsection we shed light on the importance of this "heterogeneity bias". We compare estimates of a pooled coefficient model to those of the random coefficient model. The former imposes parameter equality over all cross-sections and is predominant in the literature. The latter allows heterogeneous behaviour between banks and is the preferred model for each panel in our analysis (see Table 2). Column 9 in Table 5 presents the results. Long-term pass-through estimates are rather similar. A few exceptions notwithstanding, this is also the case for immediate pass-through estimates (at least from a statistical point of view). Similar to the aggregate results, however, we find systematic underestimation of adjustment coefficients in estimating the pooled model. The bias in the adjustment coefficient is an indication of large underlying differences between banks (see also Chen and Engel (2004)).

[18] There might be a concern that, due to the series selection described in Section 2 our data may no longer be comparable with the ECB retail rates. As a robustness check, we create mean series from our dataset and reinvestigate aggregation bias. All conclusions remain.

Illustrations of and theoretical expressions for heterogeneity bias are provided by Robertson and Symons (1992), Pesaran and Smith (1995) and Pesaran et al. (1996). Intuitively, both the aggregate and the common coefficient approach average the data before estimation, whereas the heterogeneous model averages afterwards. Averaging the data before estimation has the effect of eliminating heterogeneity. Whenever cross-sectional units are indeed heterogeneous, imposing them to behave identically will result in inconsistent estimates. The inconsistency is due to the fact that, as heterogeneity is suppressed, it will show up in the residuals. As a result, there will be a non-zero correlation between regressors and residuals, thus violating the OLS-assumptions. The results in Table 5 show that in our sample the inconsistency is most severe in the estimation of adjustment coefficients. Figure 3 visualizes the systematic underestimation in adjustment coefficients, by plotting them over the different models.

What the discussion of aggregation and heterogeneity implies for pass-through research is the following. From an economic point of view our results indicate that banks respond quicker to changing money market rates, and, by implication, to policy rates, than was previously believed. Such information is relevant in that it can help understand (lags in) the monetary transmission mechanism. One should be careful in interpreting point estimates based on aggregate series, as these are potentially misleading. Analysis of disaggregate series eliminates aggregation bias, but may still suffer from a bias due to heterogeneity. The solution to this problem lies in incorporating heterogeneity into the model. This section does so by estimating random coefficients, which also control for measurement error. The following section in addition considers systematic heterogeneity.

4. Determinants of heterogeneity in pass-through

In this section we attempt to uncover any structural determinants of the differences across banks in terms of pass-through. We investigate whether bank characteristics, such as the balance sheet structure, the risk profile or the market share of the banks, systematically cause a faster or more complete pass-through of market to retail interest rates. We start by addressing some methodological issues. Next, we give an overview of the factors that might explain cross-bank heterogeneity in adjustment. Subsequently, we estimate the model and interpret the results.

a) The estimation procedure

In the previous section we examined the pass-through of market to retail interest rates for a number of loan and deposit products. Here the objective is to investigate variation in the pass-through for the cross-section of banks, using bank-specific characteristics. Because of the limited number of banks in the sample and the fact that we do not exploit time variation, we apply the following estimation strategy. From section 3 we have both bank-specific and average (Swamy or Phillips-Moon) pass-through measures. We use this information to investigate whether bank-specific characteristics explain deviations from the average behaviour. To this extent, we regress the differences between individual bank and average pass-through coefficients on a number of bank characteristics. The characteristics we consider are structural in the sense that they capture typical features of banks that do not change very much over time, such as balance sheet structure or market position[19]. In the analysis, we pool along the product dimension, assuming that bank characteristics isomorphically affect the pass-through across products. However, we maintain the distinction between loan and deposit products, because their pricing may be driven by different factors. Thus, for each of the three pass-through measures (the immediate pass-through, the long-run pass-through and the adjustment coefficient), we pool the individual deviations from the average estimator into one dependent variable across the six loan products and across the eight deposit products. This procedure increases the number of observations per cross-sectional regression. Moreover, product-specific effects cancel out by considering deviations *vis-à-vis* the average product-specific (Swamy or Phillips-Moon) estimates. From an economic point of view, one reason for specifying the dependent variable as deviations from the average is that we no longer need to incorporate product-specific elements among the explanatory variables. Although this can yield interesting results (relating to e.g. market concentration, market rate volatility), such regressions could suffer from omitted variable bias by not taking into account, for instance, effects of outside competition, which are very hard to measure. As a result, we only focus on bank-specific differences in pass-through.

The presence of heteroscedasticity complicates estimation of the model. Within the above setup, heteroscedasticity comes into the equation in at least three ways. First, the dependent variable contains a parameter estimate that is the result of an individually estimated equation. As each of these parameters has a different variance, the homoscedasticity condition will be violated. Second, the left hand term also contains the Swamy-estimate, which is a weighted

[19] One instance in which such characteristics may change significantly over time is the case of mergers. The fact that we treat merged banks as different units before the merger (see Section 2) implies that this effect is taken into account in the analysis.

average of all individually estimated parameters within a product category. Unless all individual coefficient estimates share the same variance, there will be some observation-specific covariance between the individual and Swamy coefficients. This covariance is a second source of heteroscedasticity. Third, all the observations within a product-class have some common variance. By pooling over products, each observation will inherit some product-specific variance, again causing heteroscedasticity. For a large part, the form of the heteroscedasticity is known. Hanushek (1974) shows how to incorporate sampling error of the first kind in the model using Feasible Generalized Least Squares. With some minor modifications, we extend Hanushek's method to also capture the second and third kind of variance. In addition, following Greene (2000), we compute White-heteroscedasticity-consistent standard errors to control for possibly remaining heteroscedasticity of unknown form.

b) Bank characteristics and interest rate pass-through

In order to identify the bank characteristics that may influence the pass-through of market to retail interest rates, we start from the fact that banks are financial intermediaries; they transform deposits into loans. The objective of a bank is to maximize its profits by charging loan rates above the risk-free rate and by offering deposit rates below the (interbank) market rate. In a complete model, one would need to fully specify the supply and demand schedules for both loans and deposits. For our purpose, we borrow from the findings on the price-setting behaviour of banks, and hence their incentives to adapt their own loan or deposit rates to changing market rates, in three types of literature, i.e. studies on the role of banks in monetary transmission (Bernanke and Gertler (1995), Kashyap and Stein (2000), Van den Heuvel (2002)), studies on the determinants of bank interest margins (Saunders and Schumacher (2000)) and the structure-conduct-performance literature in banking (Berger (1995), Vander Vennet (2002), Corvoisier and Gropp (2002)). We give a short motivation for each of the characteristics and specify their expected sign in the explanation of bank interest rate pass-through.

The literature concerned with the credit channel of monetary policy transmission has stressed the importance of banks' financial structure, in particular bank capitalization and liquidity, in determining their responsiveness to monetary policy. Poorly capitalized (Kishan and Opiela (2000)) and illiquid (Kashyap and Stein (2000)) banks are hypothesised to be relatively vulnerable to monetary, and by implication, market shocks. Moreover, banks have to maintain regulatory capital against their risk-weighted assets, implying that their capacity to expand lending depends on their capital adequacy. In line with most other research, we measure the capital position of a bank as its capital-to-total-asset ratio. We expect that the capital position of a bank and both its lending and deposit interest rate pass-through will be negatively related. As a measure of liquidity we include the ratio of cash plus securities over short term deposits in the empirical analysis. Similar to equity, we expect liquidity to act as a buffer against market fluctuations, implying a negative effect on pass-through. Contrary to most analyses in the credit channel literature, we do not consider bank size as a separate characteristic. Both from a theoretical and an empirical perspective, bank size is usually considered to proxy for some (mostly conjectured) size-related characteristics, for which data are not available. We explicitly take into account the effect of these size-related factors, leaving little independent scope for bank size in the analysis.

Since a bank's price-setting behaviour is affected by its risk, we include two specific types of risk, i.e. default and interest-rate risk, and one general measure of diversification. On the lending side, the level of default risk in the loan portfolio will determine the conditions for future lending and, consequently, the reaction of the bank to changes in market conditions. A bank will only lend to riskier borrowers when it is capable of compensating expected losses by charging higher than competitive loan rates. Thus, while the effect of default risk on loan rates will be positive, we expect a negative effect on the respective pass-through estimates. In the empirical application we approximate default risk by the ratio of non-performing loans to total credit. The maturity transformation a bank performs, on the other hand, exposes the bank's profits to fluctuations in market interest rates. The extent of interest-rate risk exposure is positively related to the relative importance of interest-sensitive assets and liabilities. Similar to Weth (2002), we approximate interest rate risk by the ratio of long-term assets over liabilities. The conjectured effect of interest rate risk on pass-through is positive. A bank vulnerable to interest rate risk requires lots of hedging activities, whose terms are typically closely tied to the market. Finally, we also include a measure of diversification in the analysis as an overall indicator of the bank's riskiness. A bank that is not only active in traditional intermediation, but also in other financial services such as insurance, investment banking or asset management is assumed to be less vulnerable to shocks in the interest rate environment than specialized retail banks. As a measure of diversification, we construct the ratio of non-interest operating income to total operating income. The expectation is that changing market conditions will have only a limited effect on the loan and deposit prices of diversified banks.

Another obvious determinant of bank pricing behaviour is the degree of competition in the loan or deposit market. Since Berger and Hannan (1989), tests discerning between the structure-conduct-performance and efficiency hypotheses in explaining bank margins and bank profitability have attracted considerable interest in empirical banking. Due to the construction of our dependent variable as a deviation from group averages and the cross-sectional nature of the estimation we cannot include a general measure of the bank market structure such as a Herfindahl index. Instead we rely on bank-specific indicators of market power, i.e. the banks' market share in the loan or deposit market. The market share measure is calculated for each of the loan and deposit products. This is consistent with the relative market power hypothesis advanced by Berger (1995) which states that banks with large market shares may be able to set interest rates autonomously, especially in an oligopolistic bank market structure. We also recall that we control for the effect of product-related determinants by considering deviations from the average pass-through measure. A negative effect of the market share variable on the pass-through would corroborate the relative market power hypothesis. The alternative hypothesis is that bank pricing decisions are driven by the degree of operational efficiency as opposed to its market power. The rationale is that efficient banks will have the incentive to pass this advantage on to their customers in the form of below-average lending rates or above-average deposit rates, thereby allowing them to increase their market share. We measure the degree of each bank's operational inefficiency with the cost-income ratio and we expect a negative relationship with the estimated pass-through intensity.

Finally, similar to Gambacorta (2004), we include two variables in the loan regressions to measure possible effects of relationship lending. On the lending side, the percentage of long-term loans in total loans is intended to proxy for long-term contacts between a bank and its customers (see Berger and Udell (1992)). The hypothesis is that relationship banks will tend to smooth market shocks for their customers by smoothing interest rates over the business cycle. We also include the ratio of demand and savings deposits to total deposits to verify the thesis of Berlin and Mester (1999) who suggest that banks with a stable pool of deposits, which leaves them less vulnerable to exogenous interest rate shocks, will provide more loan interest rate smoothing.

Inclusion of the above variables also allows us to investigate the importance of some of the different channels of monetary policy transmission. The issue has proved to be somewhat involving in the literature. The first and most general distinction relates to the money and credit view. In the previous section we examined the extent of pass-through. These pass-through measures give an indication of the combined strength of transmission through all channels. In this section, we examine cross-sectional variation in responses of interest rates. As the money channel is silent on distributional effects, the search for determinants of heterogeneity will shed light only on the credit channel. Within the credit view, we focus on the two channels operating through the banking sector.

First and foremost, the bank lending channel posits that in the face of shocks, illiquid banks are unable to insulate their loan supply (Kashyap and Stein (2000)). Liquidity can be interpreted broadly and should capture at least two possible shock-offsetting mechanisms. First, banks with a substantial buffer stock of liquid assets will be more able to sustain lending policies, regardless of monetary or market movements. Second, even when such a buffer does not exist, some banks are able to fund themselves relatively easily in the deposit market[20]. The latter source of funds should be accessible especially for well-capitalized banks (Kishan and Opiela (2000)). The second subchannel of the credit view we discuss is the bank capital channel (e.g. Van den Heuvel (2001)). Via this channel, monetary (and real) shocks affect banks' loan supply by influencing the ability of banks to acquire equity in the face of a capital constraint. Monetary shocks, for instance, affect banks' profitability. A natural channel through which this operates is via the maturity mismatch in banks' balance sheets. Thus, the extent to which profits fluctuate with these shocks depends on the amount of interest rate risk a bank is exposed to. Banks are able to raise equity only when profits are sufficiently high[21]. Bank equity, in turn, determines bank lending through the existence of capital requirements.

c) Results

The results are summarized in Table 6. This table shows the estimated coefficients per equation horizontally, or per bank characteristic vertically. We also report t-statistics, the range of the bank characteristics and the implied differences in pass-through between the 25th and 75th percentile bank. Recall that 1) the left hand side is specified as deviations from the average and 2) the adjustment coefficient is negative, such that a negative effect on this measure implies a faster pass-through.

The results show that the baseline model is able to capture heterogeneity mostly for the long-run pass-through and least for the adjustment coefficient. This finding is in contrast with Gambacorta (2004), who finds no fixed heterogeneity in the long run. The rather low goodness-of-fit measures in Table 6 are not only a common finding in cross-sectional regression, but possibly an indication of random heterogeneity in the adjustment process. As Pesaran et al. (1999) suggest random heterogeneity may vanish in the long run. Our finding of

[20] This is the point Romer and Romer (1990) make for the banking sector as a whole, but with which credit view (bank lending channel in particular) proponents disagree.
[21] This holds regardless of whether equity is financed internally or externally.

considerable unexplained variation in the short run relative to the long run could be an indication of such a phenomenon. The constant is insignificant in all regressions. This is due to a combination of two factors. First, recall that the dependent variable is specified as the bank-specific coefficient minus the Swamy/Phillips-Moon coefficient. Second, in Appendix A we show that the Swamy-estimates are not that different from mean-group coefficients. Combining both elements, this implies that the left-hand side is demeaned before estimation. The fact that the constant is insignificant in the regression of the long-run pass-through suggests that the Phillips-Moon estimator, too, is not that different from the mean of the individual cointegrating coefficients.

Among the economic variables, we find no evidence confirming the relationship lending hypothesis. The long-term loans and core deposit indicators are always insignificant. The estimated adjustment coefficient, the coefficient for which the relationship lending hypothesis has the clearest prediction (see Berger and Udell (1992)), even has the wrong (negative) sign. This finding contrasts with the evidence of Berlin and Mester (1998) for the US. They find that banks with a stable pool of core deposits exhibit smoother price setting behaviour. This does, however, not necessarily preclude the existence of relationship banking in the Belgian market. Relationships can, for instance, manifest themselves in the availability of loans (credit lines), rather than in their conditions.

Turning to competition effects, we find that efficiency never induces banks to price more competitively. If operational efficiency were to strengthen the pass-through, we would expect to find a negative (positive) sign for the cost-income ratio in the regressions with the short-term or long-term pass-through (adjustment) coefficient as the dependent variable. If anything, we find the opposite. But only in the case of the adjustment coefficient for loans (-0.784) and in the case of the long-term pass-through for deposits (0.598) is the unexpected sign significant. Hence, inefficient banks are generally found to mimic market conditions more closely than the average bank. For the short-term pass-through on loans, we find that a bank with a high market share exploits its market power by following market movements less closely (the coefficient is -0.005 and significant). For the long-term pass-through, we find that the market share variable is significant at the 10% level only. However, other specifications - dropping insignificant variables[22]- confirm the significantly negative effect of market shares on the pass-through. At first sight, the economic significance, measured by the implied difference in pass-through between the 25th and 75th percentile bank, seems rather small (-0.021 for the long-term and -0.033 for the short-term pass-through). Notice, however, that the distribution of market shares is highly skewed, indicating that market power is concentrated in a few banks. The implied difference between the 25th percentile and the bank with the largest market share is as large as 34 basis points in case of the short term pass-through. Market shares never influence the adjustment speed in significant ways. In sum, for loans it seems that market power will give rise to less competitive pricing policies, strengthening the case for the relative market power hypothesis. This may have implications for the application of national competition policy in the banking sector.

Credit risk never has a significant impact. First, it could be that default risk is incorporated more into bank spreads, rather than in pass-through. Second, it is not implausible that credit risk considerations fully play their role in the determination of interest rates charged to the least creditworthy borrowers. The interest rates considered here, however, are those charged to the most creditworthy borrowers. Table 6 provides evidence that interest rate risk has a

[22] Results are available upon request.

significantly positive effect on the short-term pass-through of loan rates. This is consistent with Weth's (2002) conjecture: banks with a large interest rate risk exposure require lots of hedging activities, whose terms are market-related. In the adjustment coefficient regression for both loans and deposits, the diversification measure is highly significant. Regarding the sign of the response, Table 6 shows that a high diversification variable increases the adjustment coefficient (relative to the average). Thus, the results indicate that diversified banks are slower in adjusting retail interest rates in response to changing market conditions. Hence, banks that are active in different financial market segments are able to smooth interest rate shocks.

Capital is probably the most discerning characteristic in explaining retail bank interest rate heterogeneity. The capital variable is negative and very significant in the (long and short-term) pass-through regressions for both loans and deposits. Well-capitalized banks exhibit stickier interest rate setting behaviour. In economic terms, the 25^{th} percentile bank has a pass-through that is ten to twenty basis points higher than the 75^{th}. The results also indicate that better capitalized banks are more sluggish in adjusting loan rates to market conditions. Liquidity has a similar effect on pass-through as capital, albeit less economically and statistically significant.

These results allow some observations regarding the respective channels of monetary transmission. The fact that liquidity has a negative effect on loan pass-through (the adjustment coefficient is high relative to the mean, implying slower adjustment) is supportive of bank lending channel effects. Liquid banks react less to market movements. We also find that the pass-through is higher for poorly capitalized banks. This finding is consistent with both the bank lending and bank capital channel, which are not mutually exclusive. For the bank lending channel to cause this, a well-capitalized bank should have less problems in generating liquidity, other than from liquidating its buffer stock. In the US, the typical alternative source of liquidity is certificates of deposits. In EMU, and Belgium in particular, the use of such certificates is not widespread. The response of the liabilities in the analysis may give a more meaningful indication. Indeed, it seems that well-capitalized banks set prices less competitively. To the extent that this is due to depositors perceiving these banks as less risky[23], this is additional evidence in favour of the bank lending channel. In order to test for the bank capital channel, we interact our capital measure with the interest rate risk variable. The channel predicts that poorly capitalized banks with a large exposure to interest rate risk react more to monetary and market shocks. If the bank capital channel is at work, we should find a significantly negative interaction effect. The estimation results (Table 7) show that the interaction variable is insignificant and often has the wrong sign. Hence, we find no evidence of a bank capital channel in Belgium. Note that in Table 7 the significance of liquidity is reduced. While this somewhat weakens the evidence in favour of the bank lending channel, the capital variable is still very significant. Recall that, especially since the bank capital channel is identified separately, this still supports the bank lending hypothesis, through heterogeneity in the ability of raising deposits.

[23] Kashyap and Stein (1995) show for the US that after a monetary restriction different types of banks face a similar drop in deposit quantities. If some banks set prices less competitively, this implies that these banks have easier access to alternative deposits, probably because they are perceived as less risky by depositors.

5. Conclusion

We analyse the pass-through of market conditions to retail bank interest rates in Belgium. The dataset covers bank-level interest rates for fourteen products, both assets and liabilities, with long and short maturities and oriented to both firms and consumers. First, we measure the extent of pass-through for each product. Our modelling approach allows for measurement error and heterogeneous behaviour among banks, two factors we argue are important in working with retail interest rates. On the one hand, measurement error is incorporated because typically, bank retail interest rates are composed on the basis of inquiries, rather than being directly observed. Heterogeneity, on the other hand, deserves attention because theory does not imply all banks should respond in the same fashion. Turning to our results, we find that corporate loans adjust both quicker and more complete to changes in money market rates of comparable maturity, relative to consumer loans. This finding is possibly a key element in understanding aggregate consumption and investment responses to monetary policy. Concerning bank liabilities, market rates quickly transmit in both time deposits and savings bonds. Demand and savings deposits, however, exhibit a very sluggish response that is far from complete. Hence, banks appear to consider the segments of current and savings accounts versus time deposits and savings bond as two distinct savings markets. Within each product category, we find that the extent of pass-through is positively related to maturity. Second, the launch of EMU has generally not resulted in more competitive pricing in the banking market. In particular, for consumer loans and demand and savings deposits we find that the pass-through is more sluggish and less complete in the post 1999 era. Third, we assess the importance of four biases that may affect pass-through estimations. We find that lagged dependent variable bias and nonlinearity bias are negligible. We do find a role for aggregation and heterogeneity bias. Our results indicate, among other things, a systematic underestimation of adjustment speeds when the analysis is either based on aggregate data or when the model fails to incorporate heterogeneity. Fourth, we verify whether bank-specific determinants can explain some of the heterogeneity found in banks' interest rate setting. We find that banks with a relatively high degree of capital coverage and relatively liquid banks exhibit a more sluggish and less complete pass-through of market to retail interest rates. These results point to the existence of a bank lending channel**,** rather than a bank capital channel. Furthermore, we find that banks with large market shares set prices less competitively, supporting the structure conduct-performance hypothesis. This may have implications for the application of national competition policy in the banking sector.

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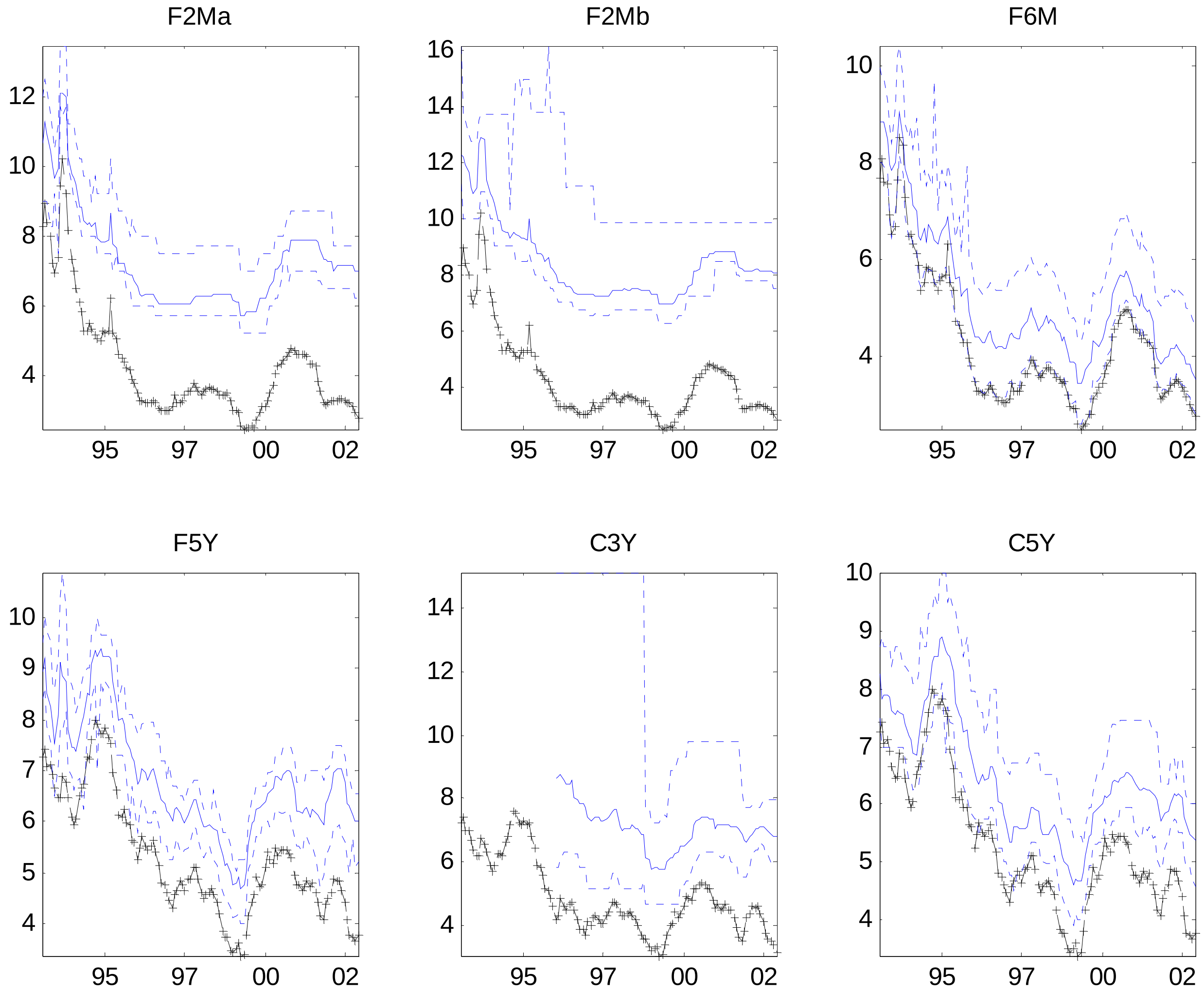
Figure 1: Summary statistics (see Table 1 for product definitions)

The figure plots the evolution of the mean product-specific interest rate (solid line) across all banks with reported rates in the sample, the evolution of the highest and the lowest (both dashed line) product-specific rate charged by any bank in a given month and the evolution of the market interest rate with the same maturity as the specific product (dotted '+' line).

LOANS



DEPOSITS

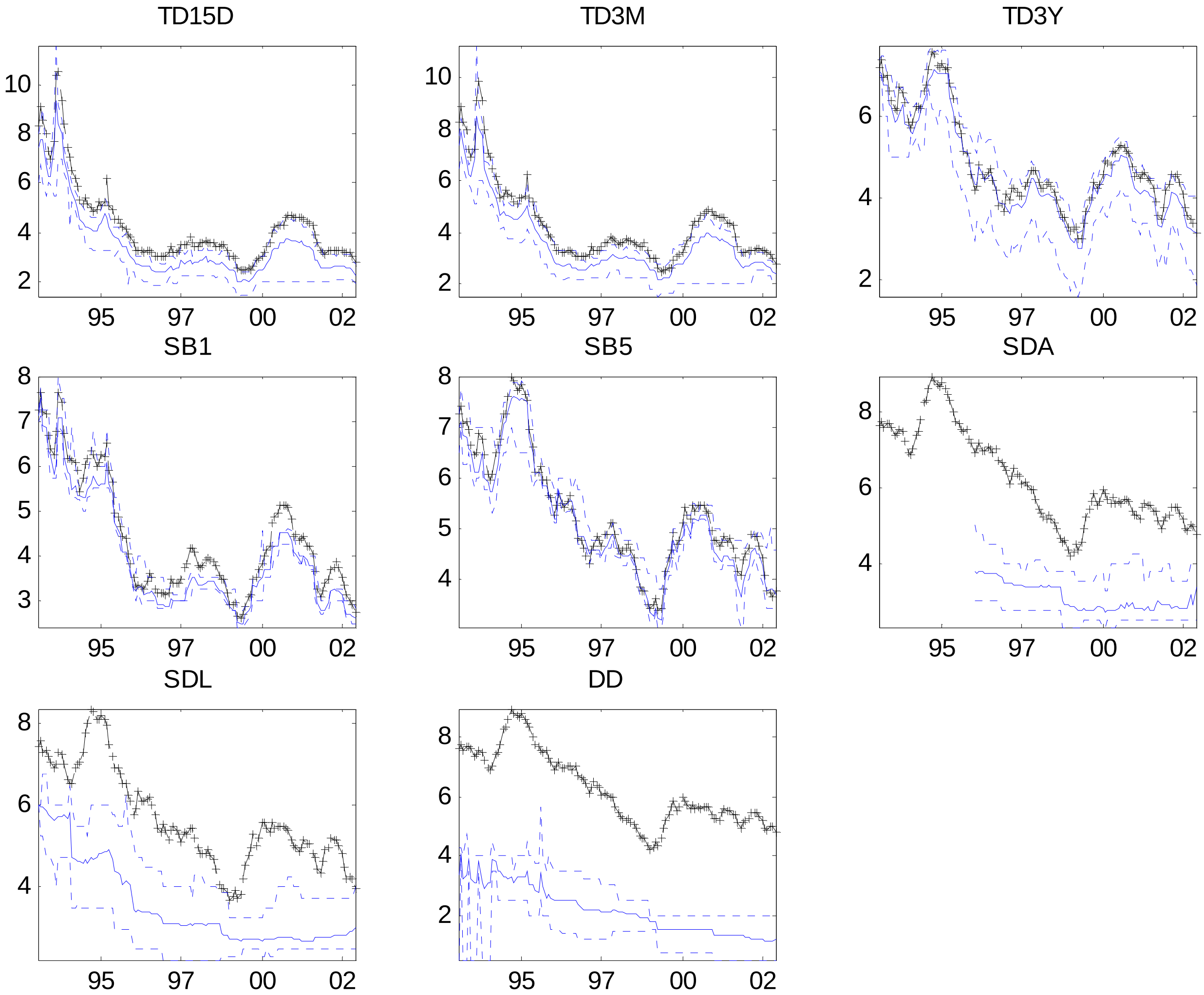
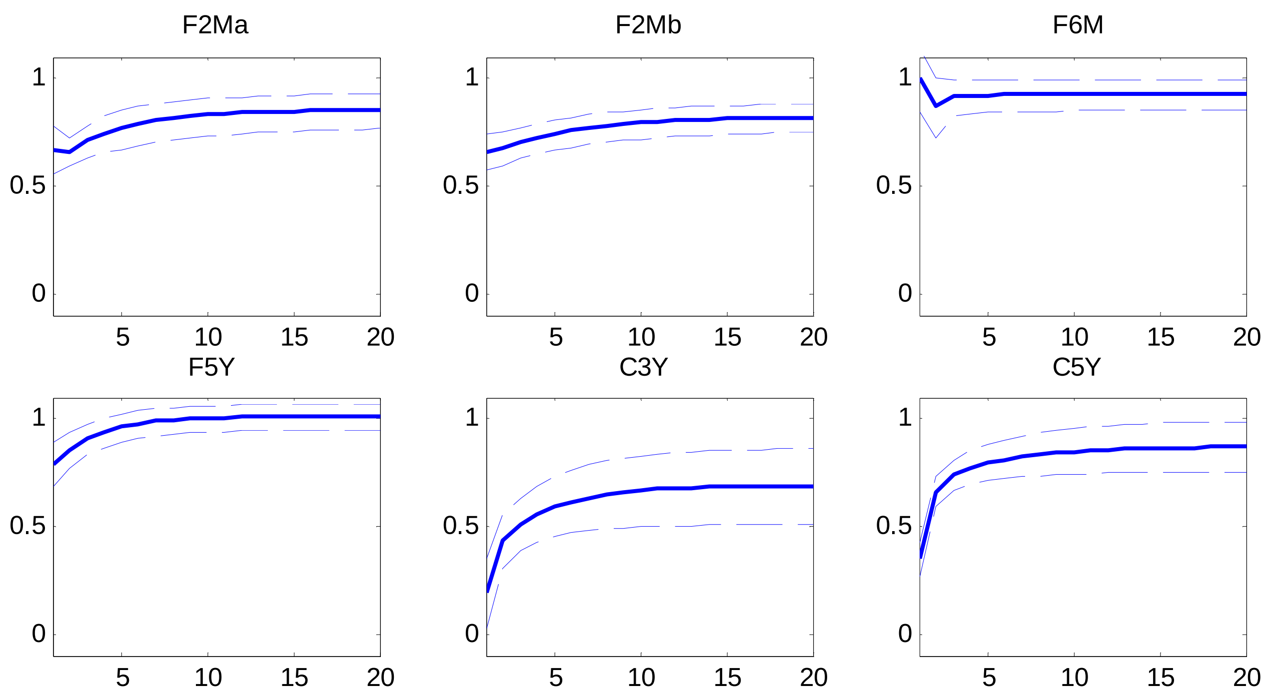


Figure 2: Pass-through: measurement (see Table 1 for product definitions)

This figure plots the pass-through for the different products. At each date (X-axis, in months) the pass-through measures the contribution of a 1%-point permanent increase in the market rate to the bank retail interest rate. Confidence intervals (95%) are computed from 5000 Monte Carlo draws, ruling out explosive roots.

LOANS



DEPOSITS

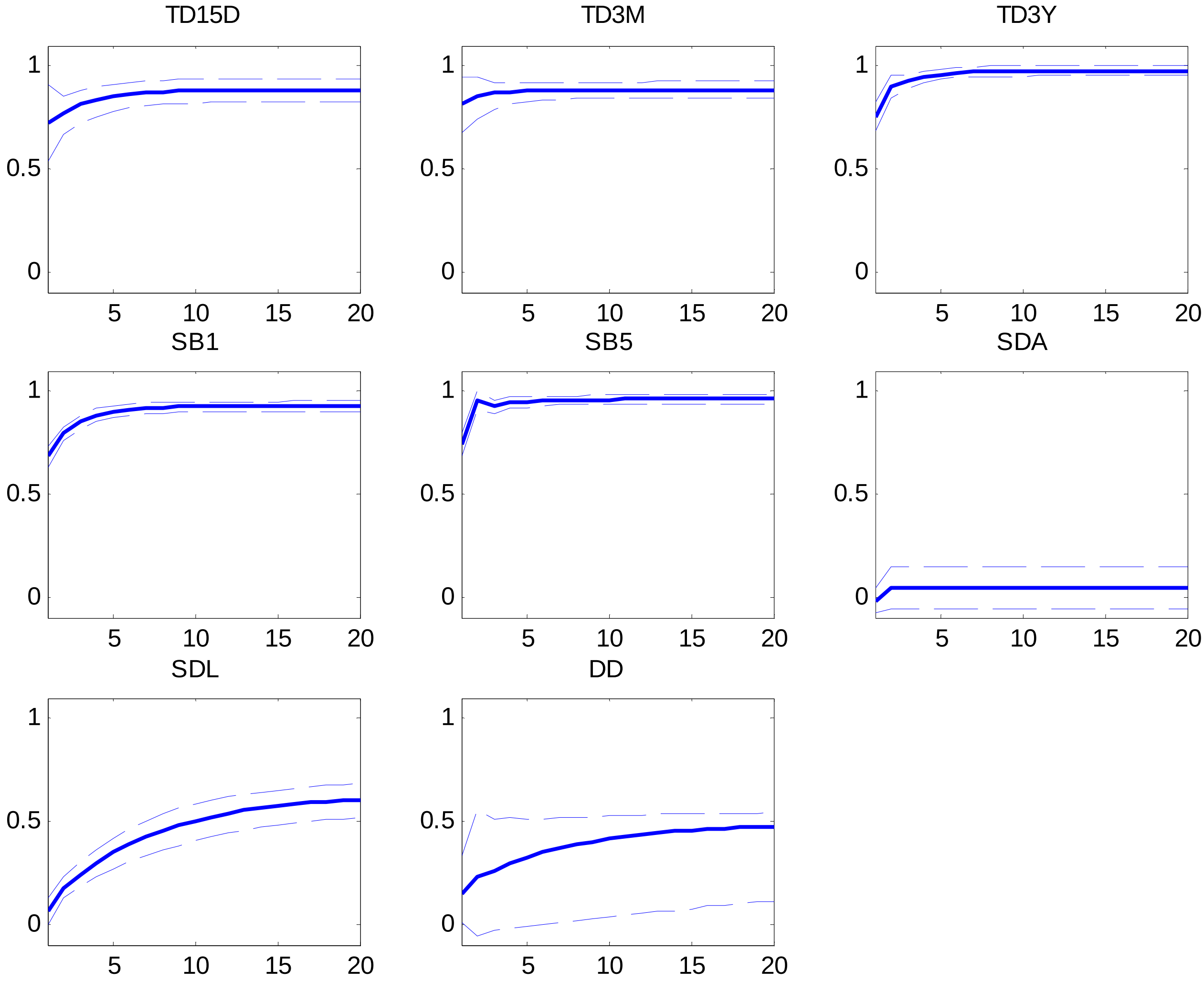


Figure 3: Adjustment coefficients

The chart plots the absolute value of the adjustment coefficient in three different pass-through models. The first ('heterogeneous') measures the pass-through based on disaggregated series and allows for heterogeneous behaviour among banks. The second ('pooled') estimates the pass-through on disaggregated series, but imposes a common reaction for all banks. The third ('aggregate') model averages interest rates into one aggregate interest rate (as reported by the ECB) and subsequently estimates the pass-through.

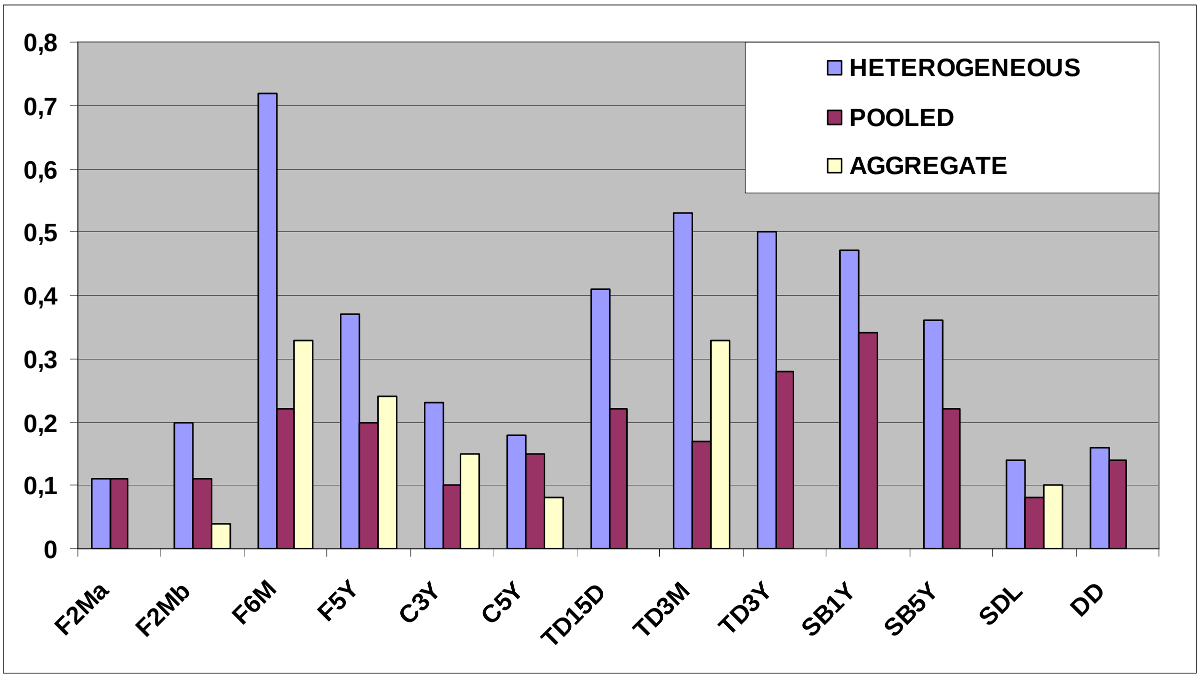
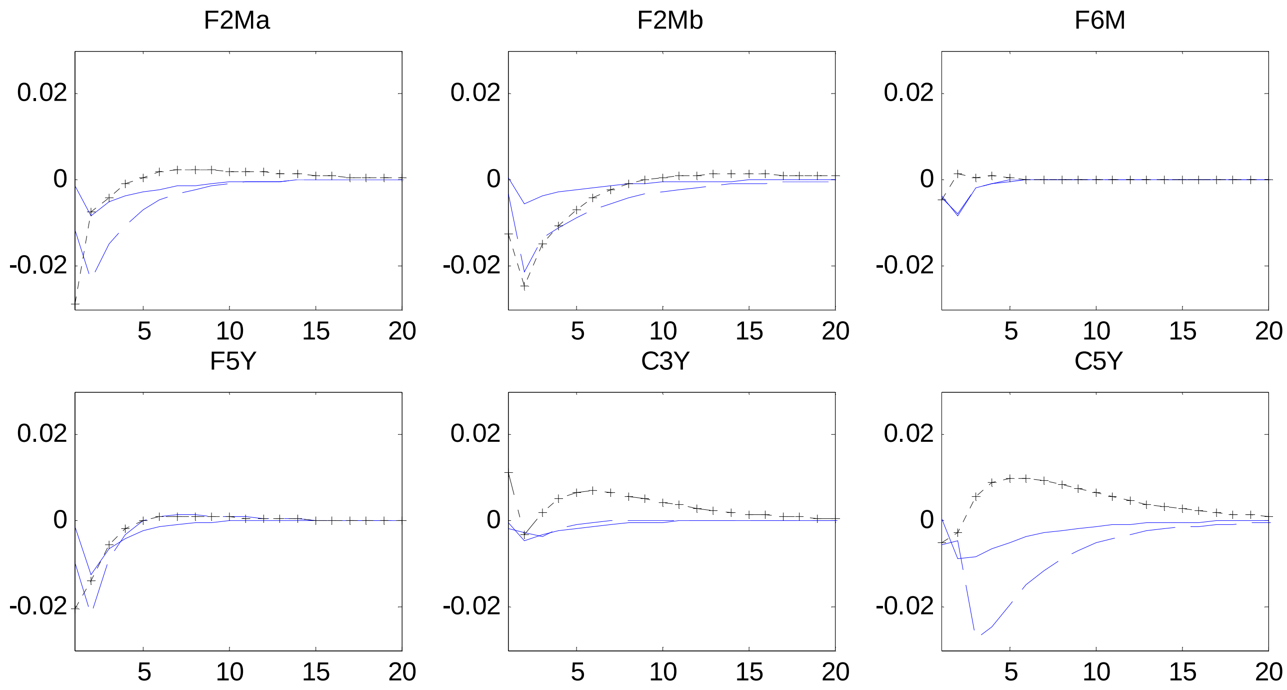


Figure 4: Mean-group (dotted '+' line), naïve bias corrected (solid line) and bootstrap corrected pass-through (broken line) expressed as deviation from the baseline model (see Table 1 for product definitions)

LOANS



DEPOSITS

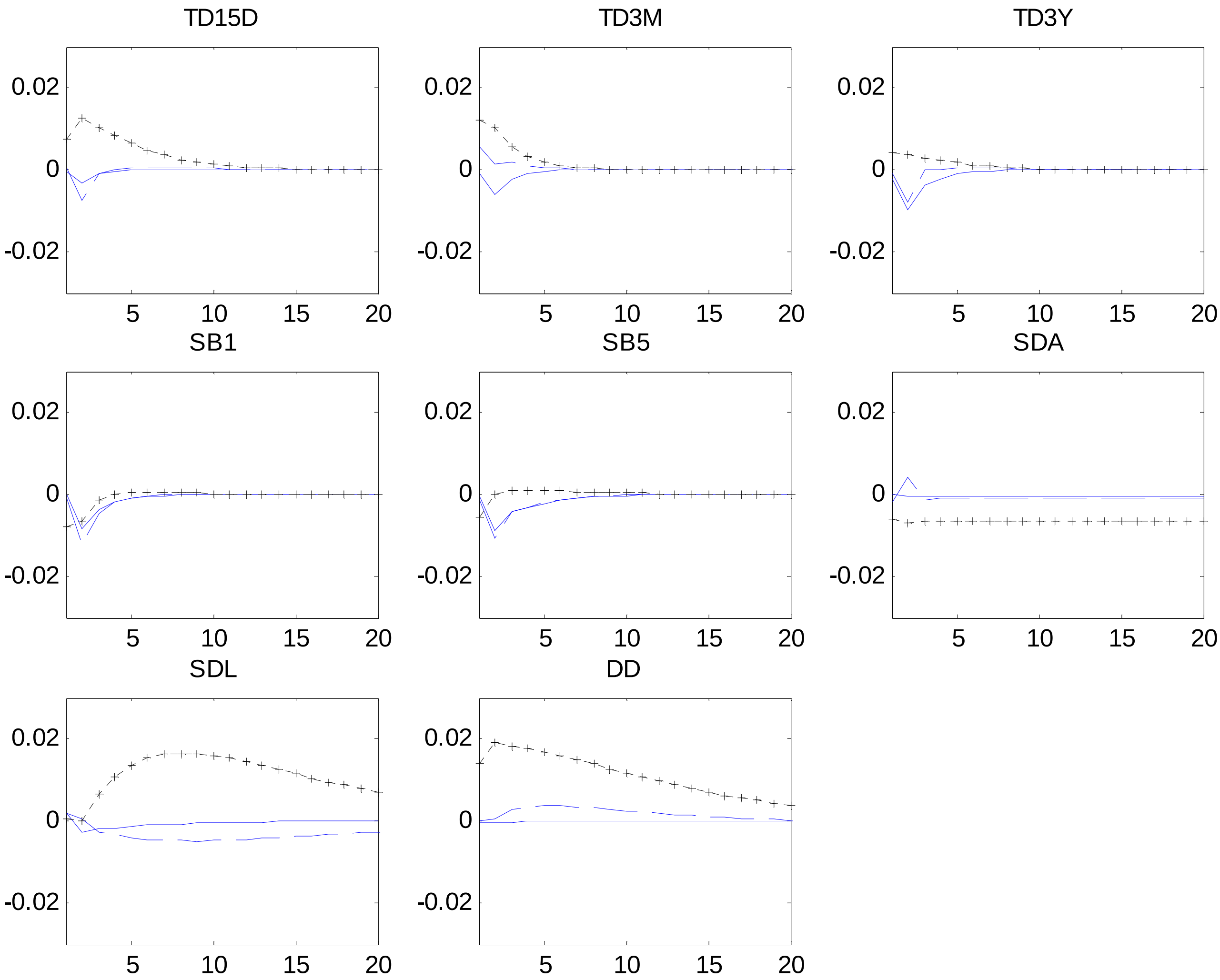


Table 1: Product definitions, abbreviations and maturity of the comparable market rates

Abbreviation	Product	Maturity
LOANS		
F2Ma	Trade certificates	2 month
F2Mb	Bank advances on current accounts	2 month
F6M	Term loan	6 month
F5Y	Investment loan	5 year
C3Y	Consumer credit	3 year
C5Y	Mortgage	5 year
DEPOSITS		
TD15D	Term deposits	1 month
TD3M	Term deposits	3 month
TD3Y	Term deposits	3 year
SB1	Savings bonds	1 year
SB5	Savings bonds	5 year
SDA	Savings deposits (+ accrual premium)	15 year
SDL	Savings deposits (+ loyalty premium)	7 year
DD	Demand deposits	15 year

Table 2: Average correlation with market rates of different maturity

LOANS	**1 mth**	**2 mth**	**3 mth**	**6 mth**	**1 y**	**3 y**	**5 y**	**7 y**	**10 y**	**15 y**
F2Ma	0,908	**0,908**	0,903	0,879	0,833	0,727	0,641	0,581	0,514	0,473
F2Mb	0,894	**0,895**	0,890	0,872	0,830	0,734	0,664	0,620	0,548	0,511
F6M	0,929	0,941	0,949	**0,961**	0,955	0,867	0,761	0,692	0,601	0,552
F5Y	0,623	0,640	0,653	0,697	0,744	0,885	**0,921**	0,913	0,878	0,841
C3Y	0,662	0,666	0,667	0,687	0,699	0,770	0,801	**0,814**	0,801	0,792
C5Y	0,666	0,679	0,687	0,721	0,754	0,849	**0,880**	0,868	0,845	0,819
DEPOSITS	**1 mth**	**2 mth**	**3 mth**	**6 mth**	**1 y**	**3 y**	**5 y**	**7 y**	**10 y**	**15 y**
TD15D	**0,910**	0,908	0,904	0,889	0,852	0,746	0,670	0,624	0,557	0,524
TD3M	0,911	0,915	**0,917**	0,912	0,886	0,784	0,701	0,651	0,572	0,533
TD3Y	0,751	0,772	0,789	0,840	0,891	**0,974**	0,965	0,929	0,874	0,830
SB1	0,920	0,936	0,946	0,969	**0,976**	0,921	0,835	0,770	0,686	0,636
SB5	0,637	0,654	0,669	0,721	0,775	0,923	**0,977**	0,976	0,958	0,933
SDA	-0,027	-0,051	-0,067	-0,084	-0,106	0,113	0,316	0,436	0,519	**0,570**
SDL	0,674	0,670	0,666	0,670	0,664	0,712	0,742	**0,772**	0,757	0,759
DD	0,489	0,486	0,484	0,497	0,505	0,577	0,628	0,664	0,684	**0,698**

Note: For each product the maturity (when specified) of the reference contract is indicated by a coloured cell.
The bold numbers indicate the maximum correlation per product.

Table 3: Likelihood ratio tests: Random versus pooled coefficient model

LOANS	LR	Critical value		DEPOSITS	LR	Critical value
F2Ma	138,79	65,17		TD15D	777,65	101,88
F2Mb	454,31	110,90		TD3M	864,82	119,87
F6M	262,49	69,83		TD3Y	222,08	101,88
F5Y	240,77	74,47		SB1	212,18	106,39
C3Y	274,73	83,68		SB5	256,76	115,39
C5Y	127,01	101,88		SDA	81,74	68,67
				SDL	238,57	110,90
				DD	77,68	26,30

Note: The table shows the likelihood ratio and the corresponding 5% Chi-square critical value. A value above the critical value indicates the restrictions of the pooled model are not valid.

Table 4: Cointegration tests

LOANS	mean(ADF)	t-stat		DEPOSITS	mean(ADF)	t-stat
F2Ma	-2,4031	-1,658		TD15D	-4,6716	-14,785
F2Mb	-2,3933	-2,1483		TD3M	-4,8777	-17,3881
F6M	-5,5363	-16,0173		TD3Y	-4,4054	-13,297
F5Y	-4,0608	-9,6107		SB1	-4,4854	-14,0676
C3Y	-3,045	-5,1239		SB5	-3,7246	-10,148
C5Y	-3,2963	-7,0992		SDA	-2,3085	-1,4616
				SDL	-2,3702	-2,0133
				DD	-2,7836	-2,0659

Note: Mean Augmented Dickey Fuller t-statistics and corrected t-statistics. The performed correction is $n^{0.5}$ * (mean(ADF)- μ)/ σ and uses μ = -2.026 and σ = 0.82 (see McCoskey and Kao (1999) for further details).

Note: reading Table 5 below

The table consists of two parts, one for loans and one for deposits. Each coloured row contains the point estimates for one (product) panel. Within a panel, we restrict attention to the main coefficients of interest, i.e. the immediate pass-through, the long-run pass-through and the adjustment coefficient. We also report mean adjustment lags. Column 3 contains the actual point estimates. The adjacent Column 4 presents standard errors on these coefficients. In Column 5 we report -in percentages- how much of these standard errors is due to parameter heterogeneity. The remaining proportion is mere parameter uncertainty. Column 6 and 7 contain mean-group and Kiviet-Phillips corrected estimates of the dynamic parameters. The former equally weighs individual cross-section coefficients, while the latter corrects for the bias due to the presence of the lagged dependent variable. Column 8 presents the point estimates of the dynamic model when applied to the aggregate retail interest rates, provided by the ECB. Column 9 contains the results based on the disaggregate series, but imposing equal dynamic behaviour over cross-sections. Finally, the right hand panel of each table contains similar information for the EMU-period subsample.

Table 5: Pass-through: measurement (see Table 1 for product definitions)

LOANS

		Full Sample (January 1993 - December 2002)							January 1999 - December 2002				
		value	standard error	hetero-geneity / total uncertainty	mean group estimator	t-bias corrected coefficient	ECB data (aggregated series)	pooled estimation (fixed effects)	value	standard error	mean group estimator	t-bias corrected coefficient	ECB data (aggregated series)
F2Ma	ST PT	0,6677	0,0634	0,82	0,6388	0,6664		0,7645	0,6256	0,2183	0,6214	0,5490	
	LT PT	0,8546	0,0379	0,75				0,9193	0,8380	0,0676			
	ADJ	-0,2660	0,0961	0,95	-0,2819	-0,2672		-0,1090	-0,1714	0,0653	-0,1749	-0,1815	
	mean lag	2,1548							3,1920				
F2Mb	ST PT	0,6588	0,0467	0,84	0,6462	0,6595	0,8104	0,6449	0,7868	0,1047	0,7852	0,7859	1,0227
	LT PT	0,8248	0,0326	0,74			0,8999	0,8706	0,8850	0,0404			0,9376
	ADJ	-0,2043	0,0629	0,93	-0,2212	-0,2056	-0,0435	-0,1082	-0,0674	0,0276	-0,0696	-0,0694	-0,0537
	mean lag	2,2122					2,5171		4,0122				3,1303
F6M	ST PT	1,0063	0,0887	0,88	1,0015	1,0028	1,0770	0,7576	1,4142	0,074	1,4141	1,4056	1,3307
	LT PT	0,9240	0,0384	0,79			0,9423	0,9296	0,9532	0,0227			0,9563
	ADJ	-0,7197	0,1527	0,92	-0,7500	-0,7276	-0,3369	-0,2208	-0,7705	0,2238	-0,7864	-0,8057	-0,9634
	mean lag	0,9822					1,4691		0,5008				0,5938
F5Y	ST PT	0,7926	0,064	0,68	0,7723	0,7914	0,8089	0,7731	0,7759	0,0987	0,7778	0,7487	0,8067
	LT PT	1,0098	0,0301	0,79			0,9213	0,9704	0,9116	0,0385			0,8541
	ADJ	-0,3727	0,0869	0,90	-0,3903	-0,3757	-0,2430	-0,2036	-0,1678	0,0813	-0,1722	-0,1884	-0,0494
	mean lag	1,6882					1,4263		1,0766				-0,7084
C3Y	ST PT	0,1906	0,0973	0,72	0,2020	0,1903	0,3302	0,2717	0,1631	0,0685	0,1684	0,1567	0,1822
	LT PT	0,6907	0,0921	0,02			0,6016	0,6881	0,5854	0,0825			0,5782
	ADJ	-0,2337	0,0514	0,89	-0,2495	-0,2342	-0,1506	-0,1014	-0,1660	0,0381	-0,1671	-0,1828	-0,2238
	mean lag	3,1337					2,3350		3,9566				2,9293
C5Y	ST PT	0,3494	0,0507	0,61	0,3444	0,3499	0,2874	0,3518	0,2440	0,0552	0,2464	0,2483	0,2301
	LT PT	0,8712	0,0602	0,96			0,8375	0,9082	0,8028	0,0487			0,8781
	ADJ	-0,1850	0,0265	0,59	-0,2042	-0,1864	-0,0839	-0,1492	-0,2942	0,0561	-0,2948	-0,3026	-0,2686
	mean lag	2,5175					2,7796		3,0484				3,6646

DEPOSITS

		Full Sample (January 1993 - December 2002)							January 1999 - December 2002				
		value	standard error	hetero-geneity / total uncertainty	mean group estimator	t-bias corrected coefficient	ECB data (aggregated series)	Pooled estimation (fixed effects)	value	standard error	mean group estimator	t-bias corrected coefficient	ECB data (aggregated series)
TD15D	ST PT	0,7200	0,0976	0,95	0,7274	0,7195		0,6513	0,9913	0,1347	1,0096	0,9928	
	LT PT	0,8844	0,0302	0,93				0,8753	0,7959	0,0897			
	ADJ	-0,4107	0,0903	0,89	-0,4408	-0,4148		-0,2233	-0,5099	0,1173	-0,5445	-0,5212	
	mean lag	1,5397							0,8864				
TD3M	ST PT	0,8124	0,0727	0,93	0,8247	0,8114	0,9971	0,7511	1,0624	0,1265	1,0601	1,0580	1,1593
	LT PT	0,8830	0,0207	0,91			0,8844	0,8783	0,7945	0,0663			0,8209
	ADJ	-0,5358	0,1136	0,93	-0,5681	-0,5413	-0,3367	-0,1717	-0,6151	0,0968	-0,6404	-0,6239	-0,1979
	mean lag	1,1494					1,2770		0,6504				0,6743
TD3Y	ST PT	0,7527	0,0407	0,80	0,7570	0,7502		0,7238	0,8812	0,0908	0,8870	0,8794	
	LT PT	0,9771	0,0128	0,61				0,9649	0,9537	0,0320			
	ADJ	-0,5044	0,07	0,88	-0,5208	-0,5089		-0,2829	-0,4699	0,1157	-0,4914	-0,4773	
	mean lag	1,4322							1,0679				
SB1Y	ST PT	0,6815	0,0286	0,70	0,6737	0,6816		0,6779	0,7571	0,0773	0,7432	0,7562	
	LT PT	0,9267	0,0127	0,87				0,9173	0,8192	0,0133			
	ADJ	-0,4736	0,0741	0,91	-0,4859	-0,4765		-0,3454	-0,3987	0,0773	-0,4156	-0,4084	
	mean lag	1,5975							1,1103				
SB5Y	ST PT	0,7393	0,034	0,77	0,7335	0,7387		0,7208	0,7979	0,072	0,7879	0,7980	
	LT PT	0,9622	0,0113	0,80				0,9578	0,9233	0,0281			
	ADJ	-0,3606	0,0582	0,86	-0,3744	-0,3652		-0,2246	-0,3448	0,0684	-0,3578	-0,3513	
	mean lag	1,3281							0,9046				
SDA	ST PT	-0,0189	0,0429	0,53	-0,0250	-0,0191			-0,0083	0,094	-0,0568	-0,0096	
	LT PT	0,0447				0,0441			-0,0592			-0,0601	
	ADJ												
	mean lag	2,4474							1,8015				
SDL	ST PT	0,0630	0,0399	0,66	0,0633	0,0647	0,0760	0,0579	-0,0509	0,0343	-0,0583	-0,0523	0,0127
	LT PT	0,6347	0,0395	0,89			0,7258	0,6474	-0,0839	0,0475			-0,0026
	ADJ	-0,1490	0,0299	0,87	-0,1636	-0,1493	-0,1025	-0,0772	-0,1279	0,0471	-0,1576	-0,1337	-0,3134
	mean lag	6,9006					11,4165		2,8916				-4,6090
DD	ST PT	0,1433	0,1031	0,86	0,1572	0,1429		0,1312	0,0390	0,0401	0,0396	0,0389	
	LT PT	0,4910	0,0333	0,42				0,5227	-0,0287	0,1067			
	ADJ	-0,1666	0,1044	0,96	-0,1776	-0,1670		-0,1430	-0,1831	0,0906	-0,1985	-0,1878	
	mean lag	5,4150							12,8615				

Table 6: Determinants of pass-through heterogeneity

LOANS

Estimation Results

		constant	Capital	Liquidity	Credit Risk	Interest Risk	Diversification	Market Share	Inefficiency	LT loans / Total loans	(Dem+Sav Dep) / Total Dep	adj R^2
ST PT	coefficient	0,233	-4.204**	0,005	-1,171	0.002**	0,357	-0.005**	-0,180	-0,020	0,154	0,151
	t-stat	0,472	-2,218	0,074	-0,937	2,333	1,408	-2,073	-0,283	-0,104	0,706	
LT PT	coefficient	-0,119	-5.452***	-0,011	1,412	0,001	0,363	-0.004*	0,049	-0,127	0,317	0,449
	t-stat	-0,219	-4,691	-0,176	1,303	1,138	1,502	-1,850	0,076	-0,828	1,634	
ADJ	coefficient	0,519	0,902	0.080*	-1,027	0,001	0.5104**	-0,002	-0.784**	-0,009	-0,004	-0,011
	t-stat	1,348	0,973	1,812	-1,582	1,450	2,384	-1,297	-1,827	-0,091	-0,025	
Summary Statistics												
	mean		0,045	1,046	0,012	9,307	0,180	5,868	0,915	0,699	0,609	
	std. dev.		0,022	0,517	0,015	22,555	0,097	10,143	0,054	0,207	0,172	
	25 percentile		0,028	0,718	0,003	2,591	0,118	0,241	0,902	0,557	0,520	
	75 percentile		0,053	1,346	0,013	4,387	0,193	6,301	0,959	0,872	0,720	
Implied differences (value at 75 percentile - value at 25 percentile)												
ST PT			-0,101	0,003	-0,012	0,003	0,027	-0,033	-0,010	-0,006	0,031	
LT PT			-0,131	-0,007	0,014	0,002	0,027	-0,021	0,003	-0,040	0,063	
ADJ			0,022	0,050	-0,010	0,002	0,038	-0,015	-0,045	-0,003	-0,001	

DEPOSITS

Estimation Results

		constant	Capital	Liquidity	Credit Risk	Interest Risk	Diversification	Market Share	Inefficiency	adj R²
ST PT	coefficient	-0,068	-2.930***	-0.095*	0,507	0,001	-0,177	0,000	0,366	0,119
	t-stat	-0,291	-3,897	-1,943	0,645	1,347	-1,311	0,101	1,318	
LT PT	coefficient	-0,282	-4.465***	-0.126***	1,094	0,000	-0,261	-0,002	0.598**	0,444
	t-stat	-1,167	-5,387	-2,883	1,182	0,184	-1,577	-0,833	2,137	
ADJ	coefficient	-0,063	2.604**	0.127***	-0,414	0,000	0.503**	-0,004	-0,297	0,039
	t-stat	-0,246	2,501	2,971	-0,463	0,235	2,013	-1,142	-1,074	
Summary Statistics										
	mean		0,050	1,039	0,014	9,092	0,170	4,452	0,907	
	std. dev.		0,024	0,506	0,017	23,210	0,089	7,633	0,068	
	25 percentile		0,031	0,718	0,002	2,383	0,111	0,128	0,870	
	75 percentile		0,068	1,331	0,014	3,848	0,190	5,894	0,961	
Implied differences (value at 75 percentile - value at 25 percentile)										
ST PT			-0,107	-0,058	0,006	0,001	-0,014	0,001	0,033	
LT PT			-0,163	-0,077	0,013	0,000	-0,021	-0,010	0,054	
ADJ			0,095	0,078	-0,005	0,000	0,040	-0,025	-0,027	

Note: reading Table 6

The table consists of two panels, one for loans and one for deposits. Each panel is composed of three parts. The first part contains the point estimates, t-statistics and a goodness-of-fit measure. Significance of point estimates at 10, 5 and 1% level is respectively denoted with *, ** and ***. The second part shows four statistics (mean, standard deviation and the value at the 25^{th} and 75^{th} percentile) for each determinant included in the specification. The last part shows the implied difference on pass-through of a change in the determinant from the 25^{th} to the 75^{th} percentile. ST PT, LT PT and ADJ stand respectively for short term pass-through, long term pass-through and the adjustment coefficient.

Table 7: Loan pass-through and the bank capital channel

Estimation Results

		constant	Capital	Liquidity	Credit Risk	Interest Risk	Diversification	Market Share	Inefficiency	LT loans / Total loans	(Dem+Sav Dep) / Total Dep	Capital * Interest Risk	adj R^2
ST PT	coefficient	0.125	-3.958**	-0.024	-0.980	0.007	0.441	-0.006**	-0.113	-0.007	0.215	-0.117	0.145
	t-stat	0.247	-1.992	-0.301	-0.725	0.504	1.399	-2.136	-0.180	-0.034	0.771	-0.390	
LT PT	coefficient	0.120	-6.085***	0.052	0.923	-0.012	0.227	-0.003	-0.095	-0.135	0.155	0.272	0.459
	t-stat	0.197	-4.290	0.614	0.765	-0.814	0.723	-1.220	-0.149	-0.877	0.515	0.901	
ADJ	coefficient	0.253	1.616	0.003	-0.324	0.020	0.693***	-0.003	-0.702*	0.028	0.239	-0.377	0.007
	t-stat	0.595	1.595	0.037	-0.398	1.550	3.002	-1.420	-1.648	0.276	0.989	-1.483	
Summary Statistics													
	mean		0.045	1.046	0.012	9.307	0.180	5.868	0.915	0.699	0.609	0.387	
	std. dev.		0.022	0.517	0.015	22.555	0.097	10.143	0.054	0.207	0.172	1.047	
	25 percentile		0.028	0.718	0.003	2.591	0.118	0.241	0.902	0.557	0.520	0.086	
	75 percentile		0.053	1.346	0.013	4.387	0.193	6.301	0.959	0.872	0.720	0.313	
Implied differences (value at 75 percentile - value at 25 percentile)													
ST PT			-0.095	-0.015	-0.010	0.013	0.033	-0.038	-0.006	-0.002	0.043	-0.026	
LT PT			-0.147	0.033	0.009	-0.022	0.017	-0.016	-0.005	-0.042	0.031	0.062	
ADJ			0.039	0.002	-0.003	0.035	0.052	-0.018	-0.040	0.009	0.048	-0.085	

APPENDIX A:

Mean-group estimates

All estimates of dynamic random coefficients referred to in the text are estimated as proposed by Swamy (1970). This subsection compares the (Swamy) measurement results with the estimated pass-through based on the mean-group estimator (Pesaran and Smith (1995)). The difference between the two estimators is the way in which individual cross-section coefficients are weighted. While the mean-group estimator simply averages (resulting in a simple to implement procedure), Swamy gives more weight to those coefficients that are estimated more precisely (improving efficiency of the estimator). We expect differences to be limited, as in composing the data we require every cross-section to exceed a minimum number of observations. This eliminates imprecisely estimated individual equations from the averaging of coefficients, and thus narrows the difference between the two estimators (having a proportionately larger effect on the mean-group estimator). As long-run coefficients are estimated as suggested by Phillips and Moon (1999), changing the weighting of individual dynamic coefficients from (1) does not alter these.

Column 6 in Table 5 and Figure 4 (dotted '+' line) summarize the results. Differences in estimated coefficients, measured by (the absolute value of) the relative deviation of the mean group point estimate from the Swamy estimate, often approach ten percent. These differences are limited in the sense that the mean group estimates never lay outside the Swamy confidence intervals. Moreover, in constructing measures of the pass-through, Figure 4 shows that hardly any differences emerge. This implies that our results are robust to minor changes in estimation strategy.

APPENDIX B:

This appendix investigates the importance of both the lagged dependent variable and the nonlinearity bias. Both these biases are theoretically plausible in a pass-through context. We here give an indication of their empirical relevance in estimation of the pass-through.

Lagged dependent variable bias

The presence of the lagged dependent variable among our regressors possibly biases the estimates of models such as (1). This bias is present in univariate autoregressive models, is exacerbated in "large n, small T" panels -where it is known as the Nickel (1981) bias- and might play a more limited role in panels having a large time dimension. In this subsection we assess whether this bias poses a problem for pass-through estimates. We implement the Kiviet and Phillips (1994) bias corrections on the estimated coefficients and recompute the pass-through.

Comparing Columns 3 and 7 in Table 5 shows that coefficients hardly change as a result of the bias correction. The bias never seems to affect a coefficient more than two percent relative to its uncorrected value. The apparent irrelevance of the lagged dependent variable bias may be due to a relatively long sample in the preceding estimations. As Kiviet and Phillips (1994) show, the bias vanishes as T increases. Although we do not conduct a full-blown analysis of the bias, trying to detect at what sample size this bias becomes important, we might get some indication in our EMU-period sub-sample estimations. These estimations are based on series with a time dimension of maximum four years. Column 13 (relative to Column 10) in Table 5 shows that although they are somewhat larger than for the whole sample, bias corrections remain limited. The reason is probably that forty-eight observations are still enough to render the bias negligible.

Nonlinearity bias

In general, the pass-through is a highly nonlinear function of estimated parameters. There are two exceptions, for which the pass-through is estimated directly: the immediate impact on the one hand, and the long-run effect in case of cointegration on the other. All other measures of transmission are possibly contaminated with a bias due to nonlinearity. Within a framework similar to ours, Pesaran and Zhao (1999) suggest several correction methods, of which we implement two. The alternative estimators correct for both the lagged dependent and the nonlinear bias.

Figure 4 plots the difference between the bias-corrected and the baseline pass-through estimates. The graph reveals that the small changes due to the lagged dependent variable bias correction hardly affect the pass-through (solid line). As Pesaran and Zhao (1999) show, this "naïve" corrected estimator performs rather poorly. They prefer a bootstrap correction, which is also shown in the figure, by the broken line. Within the Engle-Granger procedure, the long-run effect is estimated directly. Hence, there is no bias due to nonlinearity in this coefficient. This is also apparent in the figure, where the corrected estimator ultimately converges to the uncorrected estimate. Intermediate estimates exhibit the largest bias within one to six months. Overall, the bias is small. From a statistical point of view, the corrected pass-through always lies within the confidence regions of Figure 2. From an economic perspective too, the bias - showing a (absolute) maximum of less than three percentage points- does not seem to have an impact on any of the conclusions.

Overall, the previous analysis and discussion of the respective biases presents some good news for pass-through research in general. Even though there may exist some minor differences in point estimation, broad conclusions about the measurement of the pass-through do not seem to suffer from lagged dependent variable or nonlinearity bias. It should be noted, however, that these results are not necessarily extendable to research based on very short samples, nor are they guaranteed to be valid within pooled estimations.

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