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Banque de France

20 August 2008

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August 20, 2008

Abstract

This paper investigates the pass-through mechanism from market interest rates to bank interest rates using a panel of French banks based on new interest rates statistics. The data are extracted from new individual contracts, on a monthly basis for the three main sectors of the credit market (consumers loans, mortgage loans and loans to enterprises) from January 2003 to July 2007. The pass-through is estimated using recent econometric methods on non-stationary panel data. In contrast to previous studies, cross-sectional dependence among banks is allowed. Our results confirm that bank rates for loans to enterprises and mortgage loans do not adjust completely to changes in market rates, even in the long run. The model also captures the narowing of the intermediation margin during the period considered.

JEL classification: G21, C22, C23;

Keywords: transmission mechanism of monetary policy, nonstationary panel data, cross-section dependence

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1 Introduction

This paper investigates the pass-through mechanism from market interest rates to bank interest rates on a panel of 170 French banks based on new Eurosystem's harmonized Monetary Financial Institutions Interest Rates statistics (MIR). The data comes from new individual contracts, on a monthly basis, for the three main sectors of the credit market (consumers loans, mortgage loans and loans to enterprises), from January 2003 to July 2007. We will focus on the extent of the adjustment of retail interest rates in the long run following a change in market rates. Studying the pass-through with market rates of equal maturity – rather than with the policy rate – disentangle the pass-through of marginal costs and term structure effects. Banks are assumed to set their rates in accordance with the changes they face in refinancing conditions on financial markets plus a mark-up. This so-called "cost of funds" approach clearly marks the difference between the pass-through from market interest rates to retail rates and the transmission of the official rate along the yield curve.

The pass-through process is closely related to the interest rate channel which refers to the way the Central Bank can indirectly impact retail rates through the variation of bond and money market rates. However, monetary transmission also operates through a variety of complementary channels. Thus the credit channel and the bank capital channel can reinforce the effects of the interest rates channel. Through the credit channel, a rise in market rates, which hinders the collecting of deposits, affects banks' credit supply because they cannot perfectly substitute other financing sources to deposits. The bank capital channel highlights the role of bank equity capital. Therefore, as banks cannot easily raise equity, they may face a loss in capital following a rise in interest rates. Consequently, they may have to reduce credit supply or increase margins to meet capital requirements. Taking these channels into account is crucial in explaining the heterogeneity in the adjustment of lending rates across banks: the adjustment process depends a priori on banks' liability structure.

Table 1 hereafter summarizes the main findings of interest rate pass-through studies performed for the Euro area. Though these various studies differ widely in terms of scope and methods, they broadly show that a complete transmission of money market rates to bank lending rates is not achieved, even in the long run. Most of the time, the long-term pass-through is less than one.

Table 1: Pass-through estimations for the Euro area

(Adjustment of bank lending rates following 100 basis points changes in money market interest rates)

	Lending rates	
Elasticity	Short run	Long run
Mojon (2000)	0.53	.
Angeloni and Ehrmann (2003)	0.38	0.74
de Bondt (2002)	0.19-0.55	0.80-1.04
de Bondt et al. (2005)	0.04-0.38	0.48-0.74
Coffinet (2005)	0.39-0.46	0.63-0.70
Kleimeier and Sander (2006)	0.25–0.45	0.65–0.75
Sorensen and Werner (2006)	.	0.38–1.17

Note: short term = adjustment after 3 months

These studies cover the whole Euro area; the periods under review vary from one author to another

A number of studies have first examined pass-through by exploiting aggregated interest rates at the level of the country. Admittedly, the advent of the Economic and Monetary Union (EMU) and the introduction of the euro have contributed to the acceleration and the convergence of the adjustment speed across the countries of the Euro area, particularly across France, Italy and Spain. However, significant heterogeneity still characterizes the adjustment scope from one country to another. Moreover heterogeneity prevails within the same country across banking products. It is often the case that the rates adjustment for corporate loans tends to be more rapid and more complete than for households loans.

More recent papers exploit individual data at bank level and use panel data techniques. Heterogeneity can then be documented through the observation of bank individual characteristics related to their balance sheet structure. These determinants, such as the size of the bank, the amount of capital reserves or the degree of liquidity of their assets, can affect the extent and speed at which they adjust to changes in market rates. The studies based on individual data, show as well that the adjustment is higher and faster for corporate loans compared with mortgage and consumer loans. Higher pass-through is also found for credits with longer maturities. Eventually, they give evidence that heterogeneous price-setting among banks is driven by individual banks' characteristics. In Germany and Belgium (e.g., Weth, 2002 – De Graeve et al., 2004), the banks' sizes have a significant impact on the speed of the pass-through. The bigger the banks, the faster they adjust to variations in market rates. Other bank-specific balance sheet characteristics influence movements in retail rates. Thus, the higher the capitalization (capital over total assets), the liquidity (cash and securities over total assets) and the level of deposits, the stickier the adjustment (e.g., Gambacorta, 2004).

Sorensen and Werner (2006) show that heterogeneity in the pass-through is high in the Euro area using new harmonized MFI interest rates (MIR) data, available since 2003. Baugnet et al. (2007) find that Belgian banks adjust their interest rates to changes in the market rates relatively rapidly but partially, and that significant heterogeneity exists across instrument categories, sectors and maturities. However, despite the accuracy of MIR data, these latest results may be hampered by the exceptional stability of market rates during the covered period, ranging from 2003 to 2005.

The present paper provides new insights by using MIR data on a longer period, from 2003 to 2007, thus covering the recent two-year period of rising policy rates. We carry out our analysis on micro level data. We use monthly bank retail interest rates on new business that account for more than 70% of new loans granted for the period of estimation. Moreover we distinguish between three types of retail bank products i.e. corporate loans, mortgage loans and consumer loans. Finally, the present article contributes to the literature by applying econometric methods for non-stationary panel data while taking into account the issue of cross-section dependence. We focus on the long-term equilibrium relationship between bank and market interest rates and estimate it within a consistent econometric framework using the Cup-FM (Continuously Updated Fully Modified) estimator proposed by Bai, Kao and Ng (2006).

The paper is organized as follows. The article begins with some stylized facts on the evolution of interest rates and the French credit market in Section 2 and with a data description in Section 3. The econometric framework is presented in Section 4. Section 5 deals with the results of estimations and includes some robustness checks. Section 6 draws conclusions.

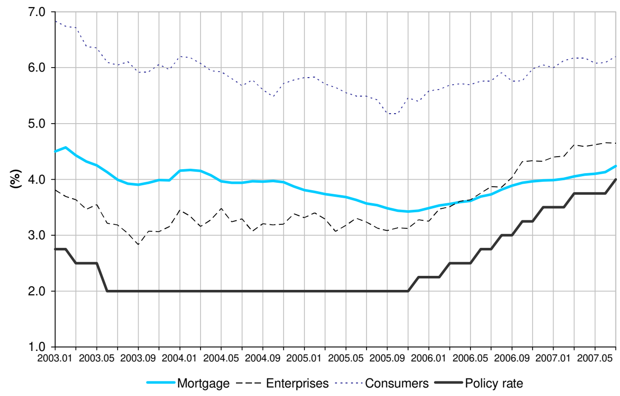
2 Stylized facts

In this section, we briefly review some stylized facts on the French credit market to allow for a better interpretation of the results.

2.1 Bank lending interest rates and market interest rates

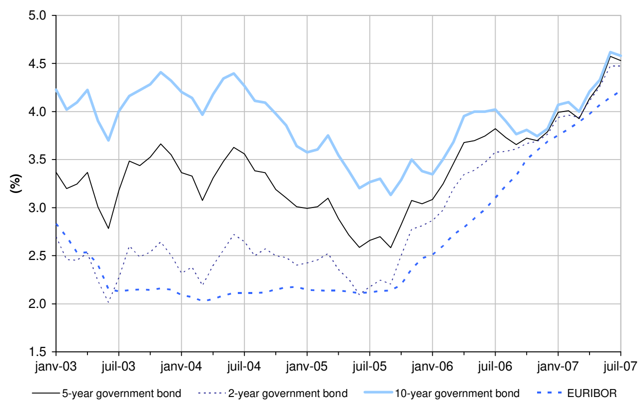
At the aggregate level, there was a decrease in the spread between the official interest rate and lending rates on new loan contracts from 2003 to 2007 (see Figure 1). According to Coffinet (2005) this could be interpreted as the result of fiercer competition in the euro area since the introduction of the single currency. Evidence shows a convergence in European retail rates for households and businesses even if there are substantial inter-country differences in interest rate levels.

Figure 1: Evolutions of the official interest rate and lending rates on new loan contracts



The beginning of the period is characterized by the pursuit of the fall in the official interest rate that was initiated on September 2001. Then, Figure 1 highlights a two and a half-year episode of stability of the official rate. Since the second half of 2005, a new cycle of rising rates has begun, with an increase in the official rate of 200 base points from November 2005 to July 2007. It is worth noting that while market rates remain quite steady between 2003 and 2005, lending rates have fallen, particularly for mortgage loans. Note that we also observe a convergence process between the rates of three types of credit over the period.

Figure 2: Evolution of money market and bond rates of different maturities



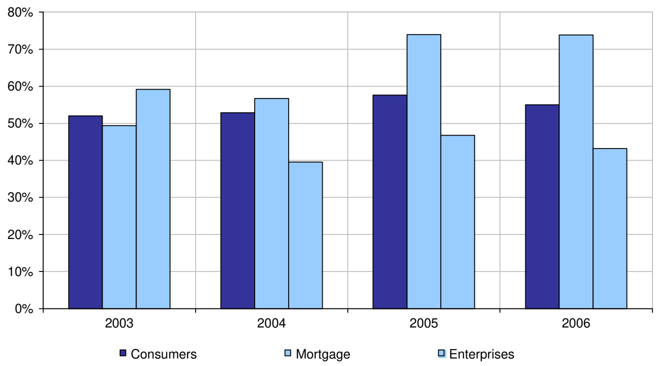
The end of the period also experiences a subsequent flattening of the yield curve as shown by the net decline in the gap between market rates, for every maturity (see Figure 2). However, from March on, market rates have risen and, in July, we notice some tightening on the money market as a result of the so-called “subprime crisis”. At the same time, a phenomenon of “flight to quality” has driven Treasury bond yields lower as investors tend to move their capital away from riskier investments. However, the results are not affected by the financial turmoil of the 2007-S2.

Under the reviewed period, loan distribution in France was particularly dynamic for both the corporate and mortgage sectors. Since 2003, the average annual growth rate of outstanding amounts for mortgage loans has been equal to 14%. After a slowdown in 2003, outstanding amounts of corporate loans have risen on average at 9% a year since 2004. Growth in consumer loans has been slightly lower. In line with this vigorous loan supply growth, lending rates have regularly decreased, up to the end of 2005.

2.2 Concentration in the French banking industry

Concentration in the French banking industry can be assessed through the percentage of operations set up by the biggest banks. As shown in Figure 3, the ten largest French banks dominate the market. They account for more than 40% of new flows of credits, on each sector market. On the mortgage loans market, their market share reaches 70%.

Figure 3: Market shares of the ten biggest French banks



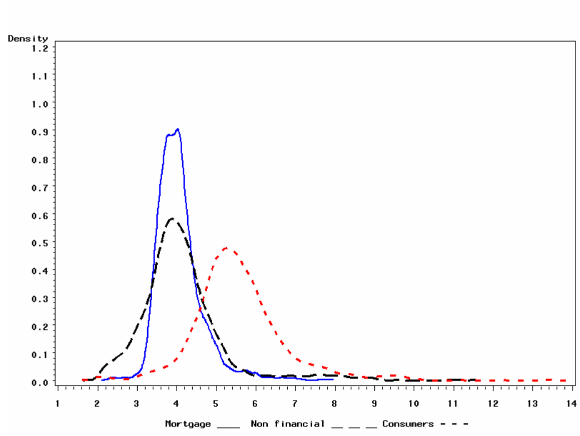
Note: market shares are computed from credit flows

Though the French credit market is concentrated, price competition is fierce. According to the results of the French Bank Lending Survey, during the last few years, competition between banks has been a major factor contributing to the easing of credit standards for loans to enterprises as well as for loans to households.

2.3 Heterogeneity in rate-setting behavior

Strong heterogeneity characterizes the rates setting across banks as well as across loans sectors. This is particularly noticeable for loans to enterprises and consumers loans as shown in Figure 4. Rates distributions are skewed for each category. They are truncated by the policy interest rate on the left, which constitutes a break even point for banks, and by usury rates on the right.

Figure 4: Density of retail interest rates (year 2004)[1]



In addition, rates settings also differ with respect to banks' size and legal category. For consumer loans and loans to enterprises, interest rates of small banks are approximately one point higher than those of large banks (See Table 2). For mortgage loans, the difference between small banks' rates and large and medium banks' rates is approximately 0.6 point. Baumel and Sevestre (2000) show that large commercial banks are more affected by competition than small banks; this is mainly due to both strong local market presence and high specialization of small banks in less competitive activities. In Section (5.2) we take into account the size effect in the estimation.

Rates settings also vary depending on banks' legal category[2]. In the sample, the most significant contrast between legal categories is found within the consumers sector, for which the difference between the mutual groups' rate and financial institutions' rate is 1.13 point. For mortgage loans, this difference is only 0.18 point. Financial institutions are characterized by a systematically higher pricing. These banks, whose scope of activities (leasing, factoring, etc) is presumably heterogeneous, make 40% of their contracts on the long-term sector, which authorises higher margins (e.g. Lacroix and Rousseau, 2007).

[1] We estimated the rates density using a non-parametric Gaussian estimator.

[2] On December 31^{st} 2005, French banks broke into: 47% of financial institutions, a third of commercial banks, 14% of mutual savings groups, the rest (approximately 6%) corresponding to the "Caisses du Crédit Municipal" (Pawn bank) and other specialized financial institutions.

Table 2: Retail interest rate (2003-2007)

Consumer loans	Mean	Std. Dev.	Median	Minimum	Maximum
Small banks	6.12	1.15	5.90	3.26	13.00
Medium banks	5.69	1.21	5.47	2.18	13.97
Large banks	5.22	0.91	5.23	3.01	8.15
Total	5.65	1.18	5.49	2.18	13.97

Loans to enterprises	Mean	Std. Dev.	Median	Minimum	Maximum
Small banks	4.78	1.32	4.57	1.64	9.71
Medium banks	4.19	1.14	4.00	1.69	11.79
Large banks	3.68	0.69	3.67	1.98	6.30
Total	4.15	1.13	3.97	1.64	11.79

Mortage Loans	Mean	Std. Dev.	Median	Minimum	Maximum
Small banks	4.58	0.81	4.43	2.96	7.96
Medium banks	4.00	0.44	3.96	2.46	6.85
Large banks	3.97	0.68	3.88	2.11	7.96
Total	4.07	0.59	3.98	2.11	7.96

3 Data description

3.1 Bank retail interest rates

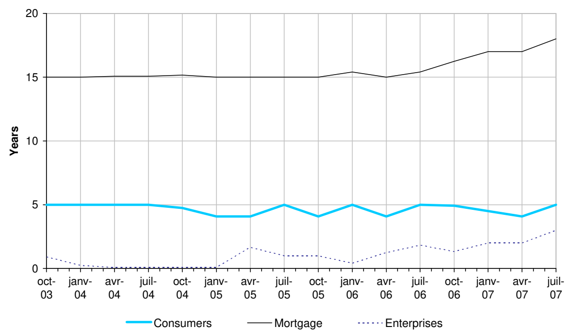
The sample includes national institutions as well as branches and subsidiaries of financial institutions operating in France but whose head office is located abroad. Data from specialized banks are collected as such, whereas data from "Generalistes" banks are extracted from a subsample of branches. Data are aggregated in order to obtain monthly estimates of flows and average rates per bank. In total, the sample includes 170 banks and covers, on average for the period, 70% of consumer loans, 88% of mortgage loans and 74% of corporate loans. After statistical treatments of the database (see Appendix A), we obtain a balanced panel of 170 banks on a monthly basis from January 2003 to July 2007. Compared to former studies carried out on French data (e.g., Baumel and Sevestre, 2000), the originality of the MIR dataset is to provide marginal rates and not average rates. These rates are applied on new contracts: consequently, they are relevant to assess pass-through insofar as they reflect conditions at the time of the agreement. It is worth noting that this corresponds to the percentage yield free of charges of the overall effective rate.

3.2 Selection of market rates

To select market rates, we adopt the so-called "cost of funds" approach which is now quite a standard approach in the literature: bank rates are assumed to be set according to the cost of banking resources. In this view, banks manage the risk of asset liability mismatch by matching the lending rate and refinancing rate for the same duration. Note that techniques of hedging and securitization are not captured in this approach. To choose the matching market rate, we refer to the median initial maturity on the considered sector of loans. We calculate this indicator using data from the Loans Survey (see Figure 5). It results in selecting a 2-year government bond rate for corporate loans, a 5-year government bond rate for consumer loans and a 10-year government bond rate for mortgage loans.

Figure 5 also shows a trend towards the lengthening of loan maturity for new contracts; it has occurred hand in hand with the rise in the policy interest rate, and has been more particularly pronounced for mortgage loans from the end of 2005. Corporate loans' maturity was very short at the beginning of the period because of the high share of loans with maturity lower than one year and has also significantly lengthened since this end of 2005.

Figure 5: Median maturity of new loan contracts



Source: Banque de France

4 Econometric methodology

Recent developments in nonstationary panel data econometrics allow us to consider the cross-section dependence betewen banks. Cross-section dependence issues are likely to arise in the credit market context, because of strategic interactions in loan pricing and spatial spillover effects.We therefore investigate cointegration taking into account cross-section dependence. Finally, we estimate the long-term relationship between market rates and retail bank interest rates using the Cup-FM estimator, as proposed by Baï, Kao and Ng (2006).

4.1 Bank interest rate pass-through model

The Monti-Klein model (see Freixas and Rochet, 1997 p. 59 for more details) provides a theoretical rationale for the behavior of banks under oligopolistic competition. In this model, the bank sets interest rates according to the following equation:

$$r^l = \frac{N\varepsilon_L}{N\varepsilon_L - 1} mc \tag{1}$$

where r^l is the retail interest rate, mc is the marginal cost of refinancing, ε_L the elasticity of loan demand and N the number of competitors. The coefficient before the marginal cost is greter than one and can decomposed in an adjustment term plus a mark-up.

In the equation (2), the marginal cost is approximated by a market rate r^m. The adjustment is complete when δ is equal to one; c is the mark-up rate which depends on the elasticity of loan demand and the numbers of competitors. In the following, the mark-up is assumed to be constant over time.

$$r^l = c + \delta r^m \tag{2}$$

We proceed in estimating equation (3) in level :

$$r^l_{i,t} = c_i + \delta r^m_t + z_{i,t} \tag{3}$$

In this approach, the bank rates depend solely on the cost of refinancing. We estimate the coefficient of the long-term relationship in the framework of panel cointegration using a panel estimator taking into account cross section dependance. But before carrying out the estimations, one has to check the properties of the panel using appropriate tests depending on the characteristics of the data at hand.

4.2 Panel unit root tests

We first investigate panel non-stationarity of the variables. Recent research in that field distinguishes between the first generation tests developed on the assumption of cross-section independence, and the second generation tests that allow, in a variety of forms and degrees, the dependence that might prevail across units in the panel. The second generation tests emerge because cross-sectional independency hypothesis is rather unrealistic in the majority of economic applications where co-movements of variables are often observed. Therefore, various tests have been proposed belonging to the so-called "second generation tests". Rather than considering correlations across units as nuisance parameters, these tests aims at exploiting these co-movements in order to define new test statistics. This approach relies on a factor structure modeling and includes, for instance, the contributions of Choi (2006) and Pesaran (2007). Here two types of panel unit root tests are employed that belong to the second generation. We implement first the Pesaran (2007) test which takes into account the existence of cross-section dependence between banks of the panel[3]. Table 3 summarizes the results for this panel unit root test. For each sector, we reject the null of stationarity. Overall, the results reach the conclusion of non-stationarity of the MIR dataset. In order to assess the robustness of our findings we also implemented the Choi test (2006).

[3] See Appendix C for a detailed description of the test based on Pesaran (2007)

Table 3: Results for Pesaran test (2007)

Model	Without fixed effects	Fixed effects	Fixed effects and trend
	CIPS*	CIPS*	CIPS*
Consumers	-2.83 (0.01)	-3.01 (0.01)	-3.53 (0.01)
Mortgage	-2.85 (0.01)	-3.06 (0,01)	-3.37 (0.01)
Enterprises	-3.35 (0.01)	-3.59 (0.01)	-4.26 (0.01)

Note: P-value in parenthesis ().

4.3 Panel cointegration tests

Several panel cointegration tests exist. The first generation panel cointegration tests as in Pedroni (1999) test for the existence of a cointegrating relationship, assuming no cross-section dependence. However, these tests have the shortcoming of not accounting for possible cross unit dependence. This, as shown by Banerjee, Marcellino and Osbat (2004) in a series of Monte Carlo simulations, leads asymptotically to substantial oversize of the tests. To overcome these problems, Banerjee and Carrion-i-Silvestre (2006) propose panel cointegration tests that model the possibility of cross-section dependence (see Appendix C). They allow for a factor structure to model dependence as in Bai and Ng (2004). In this framework, they test the null hypothesis of no-cointegration against an alternative hypothesis of cointegration using one or several unobservable common factors. The results are reported in Table 4. The null hypothesis of no cointegration is rejected: these results confirm the existence of a long run equilibrium relationship between market rates and retail rates.

Table 4: Banerjee and Carrion-i-Silvestre (2006) panel cointegration tests[4]

Model without break	H_0 : no cointegration	
	Pseudo-t $\mathrm{ADF}_{\hat{u}}\ ^c$	p-value
Consumers loans	-29.03	0.00
Mortgage loans	-42.66	0.00
loans to enterprises	-97.71	0.00

4.4 The econometric model

Oligopolistic credit markets are characterized by interdependence among banks. This dependence may be modeled through unobserved common factors which affect all the banks with different degrees[5]. Moreover, our purpose is not to estimate a model of rate-setting with all the relevant variables (competition, individual characteritics, etc.). The common factor help us to capture omitted variables. The augmented model is the following one:

$$\forall i,\ r^l_{i,t} = \alpha_i + \beta r^m_t + \lambda^{'}_i F_t + u_{i,t} \tag{4}$$

[4] We thank Anindya Banerjee and Josep Lluís Carrion-i-Silvestre for providing us Gauss codes.

[5] The dependence could also be spatial and defined with a matrix of interactions. This matrix defines, for each bank, its neighborhood formed by banks with which it is in relationship.

with F_t vector of r common factors whose one at least is I(1), and $u_{i,t}$ the idiosyncratic term. Integrated factors are not cointegrated between them and do not cointegrate with the market rate (r_t^m) . The cointegration test on panel data (see Section 4.2) is consistent with this analysis insofar as it rejects the null hypothesis in favor of a model of cointegration with at least one common factor.

We estimate the model (4) with the Cup-FM estimator proposed by Bai, Kao and Ng (2006)[6]. This estimator is asymptotically unbiased for cointegrated panel data with cross-sectional dependence. Bai, Kao and Ng (2006) propose an iterative procedure to estimate simultaneously the common factors structure $\lambda_i^{'}F_t$ and the vector β of parameters. The Cup-FM consists of two blocks: the PANIC methodology and a modified FM estimator in panel.

The PANIC methodology

We use here the two step procedure of Baï and Ng (2004) to estimate the number of common factors for the variable $r_{i,t}^l$ (see also Appendix D). The first step is to estimate the number of factors using the rates in first difference $\Delta r_{i,t}^l$. It turns out that the number of factors varies between 1 and 3 depending on the criteria used (see Table D1). The second step is to estimate among those r_0 factors the number r_1 of I(1) factors using the $MQ_c^c(m)$ tests. We cannot reject the null hypothesis of two integrated common factors (see Table D2). The robustness check based on the rates $r_{i,t}^l$ in level confirm the existence of two common trends I(1). Considering that one of the common trends corresponds to the market rate - which is I(1) - we obtain a model with only one unobservable common trend I(1) and a second stationary factor[7].

The modified FM estimator on panel data

We estimate jointly the parameters (λ_i, F_t) of the equation (5) and also the long-term pass-through β using an iterated procedure. This procedure includes a within estimator to eliminate fixed effects α_i and a "Fully-Modified" correction to account for the bias arising from endogeneity and serial correlation of the residuals so as to re-center the limiting distribution around zero (see Appendix E for a detailed description of the Cup-FM estimator).

5 Empirical Results

Estimation results for the full sample are discussed in Section 5.1, while section 5.2 checks the robustness of the results on a subsample of large banks.

5.1 Main results

We present estimation results of the long-term interest rate pass-through β in Table 6. These results show an incomplete adjustment of the retail rate to the market in the long-term. Consumer loans

[6]We implement the Cup-FM estimator on SAS V8.

[7]Estimating both factors simultaneously, results turns out to be fragile when there is a mix of I(0) and I(1) factors.

show the greater degree of transmission at 0.9 whereas the two other sectors exhibit pass-through coefficients close to 0.7. The coefficients are of the same degree of magnitude than in the individual time series approach (see Barbier de la Serre et al., 2007).

We might have expected to find interest rate pass-through equal or very close to one in the long run. The short time span of the sample could be a statistical explanation of our results. Though panel methods contribute providing efficient estimations, we would require longer series. Here, T is large in as much as our data is in monthly frequency but we only consider a little more than four years. This may be too short to fit the "theoretical long-term". However, we tend to favor the most flexible specification allowing a cointegration relationship between retail and market rates.

Economic behaviors also explain the incomplete adjustment of rates. Price stickiness theories are relevant for interest rates setting. First, banks set up long-term relationships (e.g., Berger and Udell, 1992) resulting in implicit contracts with their customers. Consequently they tend to smooth lending rates to insulate customers from market variations. Secondly because of adjustment costs when changing rates (e.g., Mizen and Hoffmann, 2004) banks adjust their rates less frequently, only when expected gains are higher than menu costs. Thirdly, limited pass-through may be due to asymmetric information between lenders and borrowers which makes banks ration credit and limit interest rates changes so as to prevent adverse selection.

Interest rate pass-through for corporate loans is traditionally higher than for mortgage loans, as large companies have access to alternative funding sources. Finally, regarding mortgage loans limited "long-term" pass-through (0.7) is linked to the specificities of the estimation period. Indeed fierce competition among banks in that market has led them to make mortgage loans a flagship product and to charge low retail rates despite the rise in market rates. Confidence intervals for panel estimations are in line with univariate analysis confidence intervals. These estimates are also consistent with those available. Baumel and Sevestre (2000) obtain an average pass-through of 0.8 with BSI data for France between 1987 and 1992.

We also verify the common factor has eliminated the cross-section dependence in the residuals. Common factors are graphed in Figure 8, with the trajectory of the average retail rate. The common factor captures the narrowing of the intermediation margin. Noticeably, the decline is sharp in 2003 for mortgage loans. Thereafter, the common factor presents a stationary trend. The decreasing trend in lending rates at the beginning of the period could be explained by stronger competition among banks. The results of the Bank Lending Survey for France also support this result. This survey shows that since 2003 the competitive pressure has simultaneously contributed to the easing of the criteria for granting loans and to a margins decrease for the majority of loans.

Table 5: Interest rate pass-through

All banks	Pass-through	Std. Dev.	Residual Std. Dev.
Loans to enterprises	0.73	0.09	0.24
Mortgage Loans	0.71	0.15	0.41
Consumers loans	0.94	0.11	0.44
Large banks			
Loans to enterprises	0.72	0.22	0.37
Mortgage Loans	0.51	0.07	0.24
Consumers loans	1.18	0.20	0.40

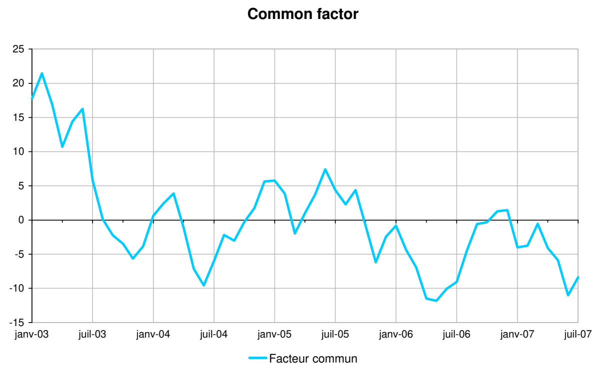
5.2 Robustness check: the size effect

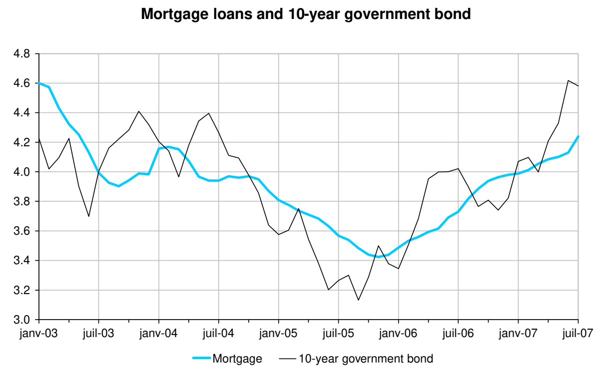
In the model (4), all the banks have the same weight when maximizing the likelihood. If the size of the bank plays a role in transmission process, the absence of weight is likely to weaken the interpretation of the slope parameters β. However, to our knowledge theoretical results do not exist concerning the weighted likelihood using panel data, we then estimate the model on the subsample of banks that belong to the six largest groups. This definition of large banks, related to the group structure, is relevant because liquidity constraints do not work in the same way for small independent banks as for small banks belonging to a group.

Compared with the results for the whole population (see Table 5 above), these new estimates are less accurate for corporate loans and consumer loans. For these two sectors, after taking into account the standard deviation, the magnitude of the pass-through estimated on the subsample of the large banks is consistent with the whole banking population's pass-through. However, for mortgage loans, banks owned by large groups show a lower long-term pass-through. Over the period, small banks react more strongly to interest rate fluctuations. This result is consistent with higher average rate registered for small banks over the period (see Table 3): small banks passed on the rising policy cycle on their retail rates since 2005 to a larger extent.

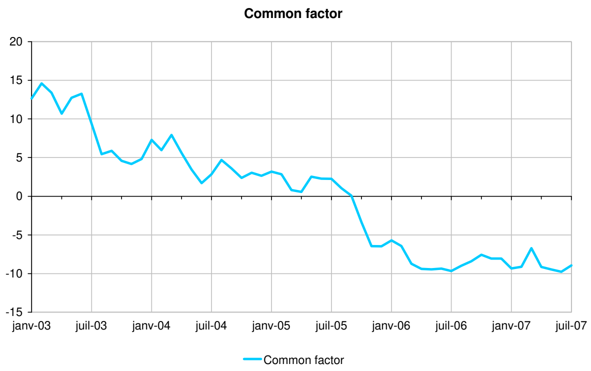
Figure 8 : Cup-FMOLS Estimations

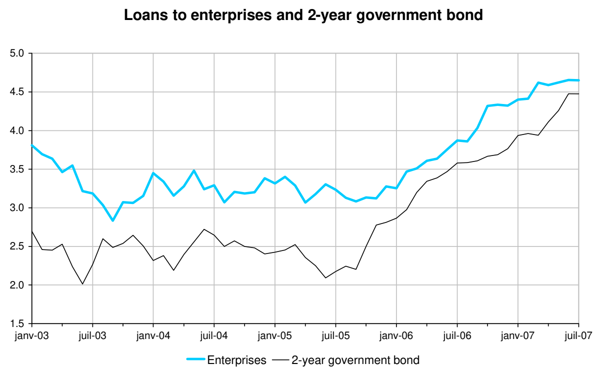
Mortgage Loans



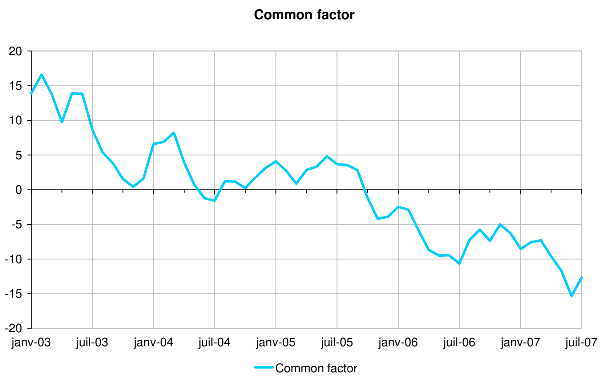


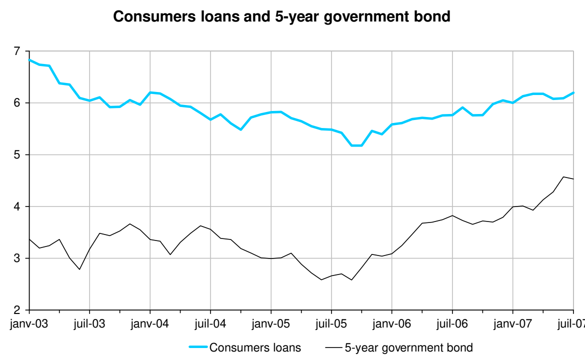
Loans to enterprises





Consumers loans





6 Concluding remarks

This paper examines the pass-through process from market rates to retail bank interest rates on a panel of 170 French banks from January 2003 to July 2007. We tackle this issue using monthly retail bank interest rates on new contracts.

First, the pass-through from market rates to retail rates is found to be incomplete and the results suggest that there is a large degree of pass-through heterogeneity across banks depending on the type of credit. Long-term pass-through coefficients range from 0.9 for consumer loans to 0.7 for loans to enterprises and mortgage loans. This incomplete pass-through reflects a rate smoothing behavior of banks. Second, if we consider cross-section dependence among banks, it turns out that a non-stationary common factor is able to explain the decreasing trend of lending rates at the beginning of the sample. We interpret this declining trend as a narrowing of the intermediation margin during the period of estimation.

This model would need additional variables. First, solvability and liquidity should be considered in order to take into account the credit channel. Second, the pricing strategies of banks also depend on monetary policy regime and on the expected policy rate. Other extensions could also be considered. One could differentiate loans following the initial maturity and fixed or variable rate, and examine how the distribution of banks portfolios between these categories changes over time. Considering a longer time span would allow analysing the issue of asymmetry of the pass-through along the rate cycle. In this context, what seems necessary is to take into account data after the financial turmoil that affected both the pricing of loans and the market rates.

7 Acknowledgements

The authors are grateful to Hervé Le Bihan for useful suggestions and discussions during an internal seminar at the Banque de France. We are grateful for suggestions from Sanvi Avouyi-Dovi, Renaud Lacroix and from Françoise Drumetz, Olivier Cousseran and Daniel Gabrielli. Lastly we thank Frédérique Savignac and Vichett Oung. For valuable suggestions we also thank the participants in the June 2008 "GDRE " in Luxembourg. The opinions expressed in this article do not necessarily reflect the views of the Banque de France.

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8 Appendix

Appendix A - Statistical treatments

The procedures of interpolation and correction of the data were performed using the Tramo-Seats programms, after eliminating the banks which had more than 16 missing observations in total over the period (see Figures below). Fusions occurred in the banks sample over the period. In order to take into account these changes of perimeter, the bank resulting from a fusion is "reconstituted": the rates and flows are calculated using the data of both establishments in the operation. We treat nine of these operations over the considered period.

Fig.1a

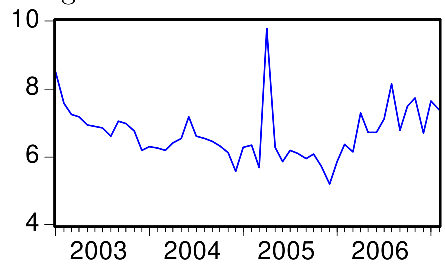


Fig.1b Once corrected

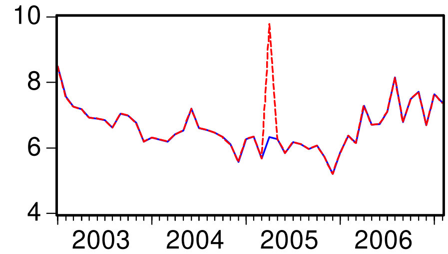


Fig.2a

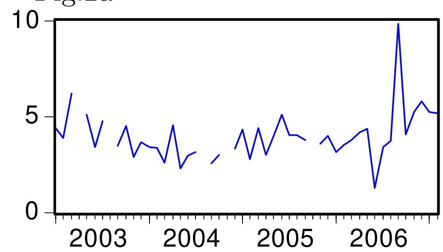


Fig.2b

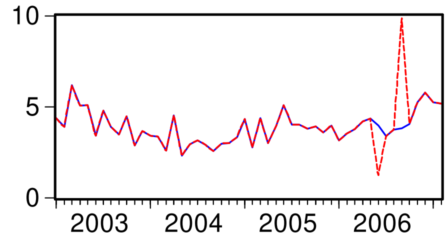


Fig.3a

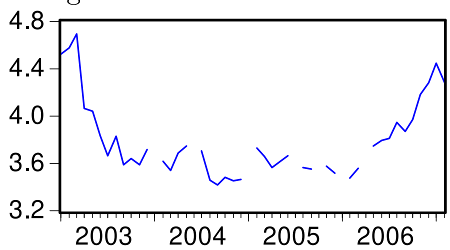
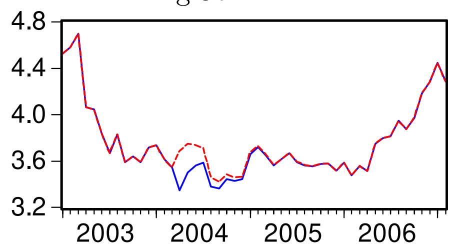


Fig.3b



Appendix B - Empirical results of the tests

Cross-section dependence in the residuals

Cross-section dependence is a strong pattern of economic series. It can arise due to omitted variable, spatial spillover effects or residual interdependence. Failure to account for cross-section dependence may invalidate estimation and inference of the long-term coefficient pass-through. So, we test for cross-section dependence between the residuals obtained for the estimations of the long-term relationship for each bank of equation 3. To do so, we use the test proposed by Pesaran (2004). It tests for error cross-section dependence and has correct size and sufficient power particularly in "large N, small T" panels. To check if the retail rates of our panel are affected by cross-section dependence, the residuals of the individual regressions of equation 3 are used to compute Pesaran's (2004) test statistic. Under the null hypothesis of no cross-section dependence, $CD \Longrightarrow N(0,1)$. The test statistic of cross-section dependence is computed as follows :

$$CD = \sqrt{\frac{2T}{N(N-1)}} \left(\sum_{i=1}^{N-1} \sum_{j=i+1}^{N} \hat{\rho}_{i,j} \right) \tag{5}$$

$\hat{\rho}_{i,j}$ are estimates of the pair-wise correlation of the residuals.

The null hypothesis of no cross-section dependence is rejected for the three sectors. Table B3 summarizes the results of the CD test A descriptive indicator is also calculated: the intraclass correlation coefficient. Similarly, the intraclass correlation coefficient is high and confirm the diagnosis. The results of CD tests indicate the presence of cross-section dependence in the residuals of the previous univariate estimates. Consequently, panel tests as well as panel estimation have to take this issue into account in order to produce reliable results.

Table B : Test of cross-section dependence Pesaran (2004) and Intraclass correlation coefficient

Long term residuals	Consumers loans	Mortgage Loans	Loans to enterprises
CD Statistics	38.19	38.94	31.24
p-values	0.00	0.00	0.00
Intraclass correlation coefficient $\hat{\gamma}$	0.36	0.30	0.52

Appendix C - Methodologyof the panel unit root tests

Panel Unit Root Test in the Presence of cross-section Dependence: Pesaran test (2007)

Pesaran (2007) considers a one-factor model with heterogeneous loading factors for residuals. Pesaran augments the standard Augmented Dickey-Fuller regressions with the cross-section average of lagged levels and first-differences of the individual series. If residuals are not serially correlated, the regression used for the ith unit is defined as:

$$\Delta y_{i,t} = \alpha_i + \rho_i y_{i,t-1} + c_i \bar{y}_{,t-1} + d_i \Delta \bar{y}_t + u_{i,t} \tag{6}$$

$$u_{i,t} = \lambda_i' F_t + \epsilon_{i,t} \tag{7}$$

where F_t is the common factor among units. Pesaran shows that the introduction of individual averages $\bar{y}_t = \frac{1}{N}\sum_{i=1}^{N} y_{it}$ and the lag $\bar{y}_{t-1}$ enables to filter the common factor F_t when $N \longrightarrow \infty$. For each unit i = *1,...,N*, we estimate this model and the statistics associated to the null hypothesis for each unit i is denoted $t_i(N,T)$. The statistics named $CIPS$, as *Cross-Sectionaly Augmented* IPS, is the cross-section averaging of the individual statistics:

$$CIPS(N,T) = \frac{1}{N}\sum_{i=1}^{N} t_i(N,T) \tag{8}$$

These statistics have similar asymptotic null distributions which do not depend on the factor loadings. But they are correlated due to the dependence on the common factor. Therefore, it is possible to build an average of individual CADF statistics, but standard central limit theorems do not apply to these CIPS or CIPS* statistics. Pesaran shows that the null asymptotic distribution of the truncated version of the CIPS statistic exists and is free of nuisance parameter and also computes simulated critical values of CIPS and CIPS* for various samples sizes. We reject the null hypothesis of unit root if the statistics $CIPS(N,T)$ is lower than the critical values tabulated in Pesaran (2007).

Panel cointegration test with cross-sectional dependence: Banerjee and Carrion-i-Silvestre (2006)

Banerjee & Carrion-i-Silvestre (2006) address the issue of cross-sectional dependence using the Bai and Ng (2004) PANIC methodology In this test, we test the null of no cointegration against the alternative hypothesis of cointegration (with up to r common factors modeling cross-section dependence).

We consider the following model:

$$y_{i,t} = x'_{i,t}\beta_{i,t} + u_{i,t} \tag{9}$$

with $i = 1, ..., N$ and $t = 1, ..., T$.The cross-section dependence is modeled by imposing a factor structure on the residuals as $u_{i,t} = \lambda'_i F_t + e_{i,t}$

We consider the following ADF regression on the estimated residuals $\hat{e}_{i,t}$ and we test the unit root hypothesis ($\rho_i = 0$)

$$\Delta\hat{e}_{i,t} = \rho_i \hat{e}_{i,t-1} + \sum_{j=1}^{k} \phi_{i,j} \Delta\hat{e}_{i,t-j} + \varepsilon_{i,t} \tag{10}$$

Afterwards, the individual ADF tests on the residuals are pooled to obtain the pseudo-t $\mathrm{ADF}_{\hat{e}}$ statistics. Specifically, they are computed as:

$$N^{-1/2} Z_{\hat{\rho}NT} = N^{-1/2} \sum_{i=1}^{N} T_{\hat{\rho}_i} \tag{11}$$

$$N^{-1/2} Z_{\hat{t}NT} = N^{-1/2} \sum_{i=1}^{N} t_{\hat{\rho}_i} \tag{12}$$

where $_{\hat{\rho}_i}$ and $t_{\hat{\rho}_i}$ are the estimated coefficient and the associated t satistics from (10)

These panel test statistics are shown to converge to standard Gaussian distribution.

$$N^{-1/2} Z_{\hat{t}NT} - \Theta_2 \sqrt{N} \Rightarrow N(0, \Psi_2) \tag{13}$$

Where the moments Θ_2 and Ψ_2 are the same as the ones for the statitics in Bai and Ng (2004).

Appendix D - Results of the PANIC methodology

First step: Selecting the number of common factors F_t

The purpose is to estimate the number of common factors for the variable $r^l_{i,t}$. We estimate the number of factors (stationary or non-stationary) using the rates in first difference $\Delta r^l_{i,t}$. Then we estimate among these r_0 factors the number of I(1) factors. In order to determine the number of estimated factors, we have computed the seven criteria suggested by Baï and Ng (2002). The number r of factors is estimated by minimizing information criteria. Table D1 below shows the number of common factors for each category.

Table D1: Information criteria on rates in first differences

Loan category	IC1	IC2	IC3	PC1	PC2	PC3	BIC3
Consumers loans	2	2	3	3	2	3	1
Mortgage Loans	3	3	3	3	3	3	2
Loans to enterprises	3	1	3	3	2	3	1

Second step: Analyzing the properties of common factors

To analyze the properties of the r_0 extracted common factors, we compute the statistics $MQ^c_c(m)$[8] developed by Stock and Watson (1988) for testing F_t along with the critical values given in Bai and Ng (2004). We want to identify the number of common trends in the vector of common factors. Table D2 reports the statistics $MQ^c_c(m)$ for each value of r_1. The results cannot reject the null hypothesis of two integrated common factors.

Table D2: Properties of common factors

r	Consumers	Mortgage	Enterprises	1%	5%	10%
1	-48.07	-39.76	-53.50	-20.15	-13.73	-11.02
2	-16.67	-24.39	-30.12	-31.62	-23.53	-19.92

Note: H_0 r = m integrated common factors

As a robustness check, we directly estimate the number of common stochastic factors using the retail rates variables $r^l_{i,t}$ in level. The methodology is similar to the previous one and is based on modified criteria IPC1, IPC2 and IPC3. The existence of two common trends I(1) is also confirmed by these criteria

[8] $MQ_f(m)$ filters the factors F_t under the assumption that they can be represented as a finite order VAR(p) process.

Appendix E - Fully Modified Cup Estimator[9]

This estimation method of panel cointegrated models with cross-sectional dependence which is modeled through the use of a factor structure[10]. We consider the model :

$$y_{i,t} = x'_{i,t}\beta + e_{i,t} \tag{14}$$

with $i = 1...N$ and $t = 1...T$

$$x_{i,t} = x_{i,t-1} + \varepsilon_{i,t} \tag{15}$$

x_{it} is a set of k non-stationary regressors and β is a $k\times1$ vector of the common slope parameters. The cross-section dependence is modeled by imposing a factor structure on the residuals $e_{i,t} = \lambda'_i F_t + u_{i,t}$.

where F_t is a r×1 vector of common factors as $F_t = F_{t-1} + \eta_t$, λ'_i is a r×1 vector of factor loadings and $u_{i,t}$ the idiosyncratic component associated to the unit i at date t. We estimates the factor parameters (λ_i, F_t) and the slope parameters β simultaneously. The estimator takes into account the fact that the "explanatory" variable β does not depend on i, eliminating the strategy proposed by Westerlund (2007)[11]. The procedure includes the Within estimator to eliminate fixed effects α_i, the Fully-Modified correction to treat long-term endogeneity, and an iterative approach to estimate all the parameters. Like the FM estimator of Phillips and Hansen (1990), the corrections are made to remove serial correlation and endogeneity.

The CupFM estimator for $(\hat{\beta},\, \hat{F})$ is obtained by iteratively solving (16) and (17):

$$\hat{\beta}_{CUPFM} = \hat{y}^+_{i,t}\left(\sum_{i=1}^{N} x'_i M_{\hat{F}} x_i\right)^{-1} \sum_{i=1}^{N}\left(\left(x'_i M_{\hat{F}}\hat{y}^+_i\right) - T\left(\hat{\Delta}^+_{\varepsilon,u,i} + \hat{\delta}'_i \hat{\Delta}^+_{\eta u}\right)\right) \tag{16}$$

$$\hat{F}V_{NT} = \frac{1}{NT^2}\sum_{i=1}^{N}\left(y_{i,t} - x_i\hat{\beta}_{CUPFM}\right)\left(y_{it} - x_i\hat{\beta}_{CUPFM}\right)'\hat{F} \tag{17}$$

The Cup-FM estimator (Continuously Updated Fully Modified) proposed by Bai, Kao and Ng (2006) results from concentrated likelihood maximization.under the identification assumption of the standard Factor I(1):

$$\frac{\sum_{t=1}^{T} F_t^2}{T^2} = 1$$

[9] Results presented here are from Baï, Kao & Ng (2006).

[10] The standard least squares estimator is, in general, inconsistent owing to the spuriousness induced by the unobservable I(1) trends.

[11] Westerlund, J. (2007) "Estimating Cointegrated Panels with Common Factors and the Forward Rate Unbiasedness Hypothesis" *Journal of Financial Econometrics* 5: 491-522.

The transition from iteration J to iteration $J+1$ is as follows:

$$\begin{cases} \widehat{\lambda}_i^{(J+1)} = \lambda_i\left(\beta^{(J)}\right) \\ \widehat{F}_t^{(J+1)} = F_t\left(\beta^{(J)}\right) \\ \widehat{\beta}^{(J+1)} = \beta\left(\lambda_i^{(J)}, F_t^{(J)}; i=1,\ldots,N \text{ et } t=1,\ldots,T\right). \end{cases} \tag{18}$$

with λ_i, F_t and $\beta(.)$ explicit functions that depend from $\left(r_{i,t}^d\right)$ and (r_t^m).

For each iteration we also calculate the residual variance of $\dot{u}_{i,t} = u_{i,t} - u_{i,t-1}$

$$\widehat{\sigma}^{(J)2} = \frac{1}{NT}\sum_{n=1}^{N}\sum_{t=1}^{T}\dot{u}_{i,t}^2 \tag{19}$$

The convergence is achieved when:

$$\max\left\{\left|\widehat{\sigma}^{(J+1)2} - \widehat{\sigma}^{(J)2}\right|, \left|\beta^{(J+1)} - \beta^{(J)}\right|\right\} < \varepsilon = 10^{-10}$$

The speed of convergence of estimators is $T\sqrt{N}$ when N and $T \to +\infty$. They are asymptotically Gaussian under the additional assumption that $N/T \to 0$. As our data rather verify "large" N in relation to T, the standard deviations must be interpreted with caution. The correction for endogeneity and serial correlation is made during each iteration. Bai, Kao and Ng (2006) show that the CupFM estimator has good finite sample properties. Moreover, this estimator is asymptotically unbiased and normal with $(N,T) \longrightarrow \infty$

$$\sqrt{N}T\left(\hat{\beta}_{CUPFM} - \beta^0\right) \xrightarrow{d} N(0,\Sigma) \tag{20}$$