IS THERE LENDING RATE STICKINESS IN THE CHILEAN BANKING INDUSTRY?

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This paper studies the transmission of monetary policy in terms of the interest rate pass-through in the case of Chile. Specifically, we are interested in the response of the commercial bank lending rate to a money market interest rate movement. International evidence suggests that lending interest rates are somewhat sluggish to adjust to changes in the policy rate. This stickiness is generally related to lack of competition in the banking sector, capital flow restrictions, and volatility of the policy rate.

One of the first comprehensive empirical studies on bank interest rate pass-through for monetary policy is Cottarelli and Kourelis (1994). They find important differences among countries: the estimated impact effects vary between 0.06 and 0.83, and the long-run effects range from 0.59 to 1.48, with an average of 0.97. Our estimates for the Chilean case are an impact of 0.81 and a long-run pass-through of 0.97 for nominal interest rates.

Previous studies suggest that sluggish adjustment is associated with market conditions and regulation of the banking sector. In this paper, we use bank-level data to explore other factors that may influence the degree of delay in market interest rate response to changes in the policy rate. The aim is to identify which characteristics may

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explain the differences in the average rates charged by each bank and their responsiveness to movements in the policy rate. The main variables considered are bank size, type of customers, and the loan risk level, which is related to demand elasticity and the cost of adjustment for banks. The theoretical model presented in the paper motivates the choice of these factors, and dynamic panel data estimation supports the implications of the model.

The paper proceeds as follows. In section 1, we briefly review the previous literature and present our own estimations for the Chilean case, at an aggregate level. Section 2 discusses some stylized facts for the Chilean banking industry and presents a model of monopolistic competition with asymmetric information for bank lending rates, together with the panel data econometric analysis. In section 3, we summarize and present some concluding remarks.

1. Chile versus the International Evidence

This section offers a brief review of empirical studies related to the flexibility of the bank lending rate in different countries. We also present our own estimations for Chile and compare them with results for other countries.

The lending rate stickiness refers to the small response of commercial banks' lending rate to a money market interest rate movement. Berger and Hannan (1989), Hannan and Berger (1991), and Cottarelli and Kourelis (1994) provide arguments and evidence for a sluggish adjustment of the lending interest rate in the short run. They find that in the long run, the lending rate fully adjusts to the shift in the money market rate. Many subsequent papers test the monetary policy transmission for specific countries under different periods and types of regulation. All of them are based on different parameterization of the following basic model:

$$i_t = \delta + \sum_{j=1}^{m} \beta_j i_{t-j} + \sum_{k=0}^{n} \alpha_k m_{t-k} + \sum_{l=0}^{p} \gamma_l \Delta \mathrm{MPR}_{t-l} \,, \qquad (1)$$

where i represents the bank-lending rate, m is the money market or interbank rate, and ΔMPR is the change in the monetary policy interest rate. The difference between the money market or interbank rate and the monetary policy rate is that the first two are interest rates determined in the market, while the latter is set by the Central Bank as a target value. In Chile, as in many other countries, monetary policy is conducted by managing liquidity, such that the interbank or money market rate is in line with the policy rate. We can therefore separate the effect of monetary policy into two steps: from policy rate to money market rate and from money market rate to lending rate; we are interested in the second step. The coefficient of interest is α_0, which indicates the impact or the short-run effect of the money market or interbank rate on the lending rate. It is expected to be positive and less than or equal to one. The coefficient that measures the long-run effect of the money market rate on the lending rate is estimated as

$$\lambda = \frac{\sum \alpha_k}{1 - \sum \beta_j} \,. \tag{2}$$

This coefficient is expected to be positive and close to one in an industry that is highly competitive.

1.1 Literature Review

In the empirical literature we find two types of studies, those that analyze monetary transmission mechanisms using cross-country data and those that give evidence using time series data for specific countries. The first group computes impact and long-run effects for different countries and then relates their findings with financial structures and macroeconomic variables of the different economies included in the sample. The second group uses country case studies to look for changes in the monetary policy transmission over time and for variation in interest rates. The main idea of both types of studies is to capture the effect of institutional features on the transmission of monetary policy.

One of the first comprehensive empirical studies on interest rate pass-through for monetary policy is Cottarelli and Kourelis (1994). This study estimates equation (1) for thirty-one countries, including developed and developing countries. They find important differences across countries in the impact coefficient, but the long-run coefficient tends to one in most cases. In a second step, they correlate the different coefficients with possible explanatory variables. The main finding here is that the impact coefficient is highly correlated with the structure of the financial system. Specifically, the lending interest rate becomes more flexible when the barriers to entry to the banking industry are low, the share of private ownership in the banking system is high, there are no constraints on international capital movement, and there is a market for negotiable short-term instruments. Neither market concentration nor the existence of a market for instruments issued by firms affects the degree of interest rate stickiness.

An important policy implication obtained by Cottarelli and Kourelis is the relevance of the discount rate or monetary policy rate as a policy instrument. In general, they argue that the movement in the discount rate is interpreted as a signal that helps reduce the degree of stickiness, especially in those economies with a weak financial structure.

Borio and Fritz (1995) examine the relationship between the monetary policy rate and the bank lending rate for a group of member countries of the Organization for Economic Cooperation and Development (OECD). Canada, Great Britain, and the Netherlands show a high short-run coefficient (above 0.7), while Germany, Italy, Japan, and Spain exhibit the highest degree of interest rate stickiness. The pass-through is more homogenous across countries in the long run, and it moves closer to one. Borio and Fritz argue that the difference in the results across countries may have to do with the type of lending rate available. In fact, interest rates for prime customers tend to adjust faster than other interest rates.

Mojon (2000) analyzes monetary policy transmission across euro area countries. He also looks for the implications of different financial structures for the stickiness of the retail interest rate. Like Cottarelli and Kourelis, he finds large differences in the short-run coefficients for different countries, ranging from 0.5 in Italy to 0.99 in Netherlands.[1] The pass-through coefficient is lower the higher is the volatility of the money market rate and the lower is the competition from other sources of finance (the level of banking disintermediation). Competition among banks reduces asymmetries through the interest rate cycle; that is, the size of the pass-through coefficient is less affected for upward movement in the interest rate than for downward movement.

A second group of studies concentrates their analysis on specific country cases. Following Cottarelli and Kourelis (1994), Cottarelli, Ferri, and Generale (1995) explore why the transmission of the monetary policy rate is so slow in Italy. They find that the high degree of

1. Toolsema, Sturm, and de Haan (2001) find similar results for the same group of countries.

stickiness is explained by the constraints to competition in the banking and financial system. Banks that operate in more competitive markets tend to translate movements of the money market rate into lending interest rates faster than do banks operating in a less competitive environment. This conclusion is based not only on the international comparison of Italian banking industry with the rest of the countries, but also on data analysis at the individual bank level. The stickiness of lending rates tends to decline with financial liberalization in Italy, which is consistent with the results using microeconomic data for different banks and regions of that country.

Using the same methodology as earlier studies, Moazzami (1999) confirms that interest rate stickiness in the United States was higher than in Canada during the 1970s and 1980s. The degree of flexibility has changed for both countries, however, moving in opposite directions over the first half of the 1990s. The short-run pass-through has thus converged to around 0.40 for both Canada and United States. The author attributes these changes to a more competitive environment for the U.S. banking system and a less competitive one for Canada.

Winker (1999) combines an adverse selection model with a marginal-cost pricing model to find an empirical equation in which the lending and deposit rates depend on the money market rate in the long run but not in the short run owing to the adverse selection problem. Based on the same argument, he justifies the lending rate's lower speed of adjustment toward its long-run level compared with the deposit rate, since the short-run coefficient for the lending rate is much smaller than that of the deposit interest rate. Winker provides evidence for his model for the case of Germany.

For the case of Spain, Manzano and Galmés (1996) use an interesting database that allows them analyze the speed of interest rate adjustment by type of bank. They define four groups of financial institutions: national banks specialized in commercial banking, savings banks, foreign banks, and merchant banks. The degree of short-run interest rate response to changes in the interbank rate varies greatly across groups, from 0.25 to 0.75 in the short-term impact coefficient. In the long run, all but saving banks have a total impact coefficient greater than one based on the reported confidence interval. In the case of savings banks, the coefficient is strictly less than one, although the deposit rate shows a higher degree of stickiness in both the short run and the long run. The impact coefficient ranges from 0.2 to 0.46, and the total impact varies between 0.63 and 0.81.

Table 1 summarizes the results of the literature reviewed.

Table 1. Summary of Results of Reviewed Literature

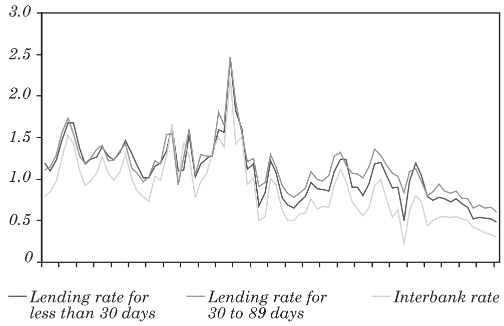
Study and sample	*Degree of transmission*	*Main conclusions*
Cross-country studies		
Cottarelli and Kourelis(1994) Sample: 31 countries	Short term: 0.06 to 0.83 Long term: 0.59 to 1.48, with an average equal to 0.97	The degree of flexibility increases with the elimination of capital flow restrictions, lower barriers to competition, private property in the banking industry, and the existence of short-run instruments
Borio and Fritz (1995) Sample: 12 OECD countries	Response to a simultaneous change in policy and money market rate Short term: 0.0 to 1.08 Long term: 0.74 to 1.17	The type of lending interest rate used could explain the differences across countries. For some countries the lending rate is applied to the best larger customer while for others the rates correspond to retail banking.
Mojon (2002) Sample: Panel data on 6 European countries	Short term: 0.5 (Italy) to 0.99 (Netherlands) Long term: Around 1 for all countries	The flexibility of interest rate increases with lower volatility of the monetary policy interest rate, and higher external and within-industry competition
Country case studies		
Cottarelli, Ferri, and Generale (1995) Italy	Short term: 0.07 Long term: 0.92	The degree of stickiness is inversely related to the degree of competition and financial liberalization
Moazzami, B. (1999) Canada and United States	Short term (CAN): 0.46 to 1.1 Short term (USA): 0.25 to 0.6 Long term (CAN): 0.6 to 2.0 Long term (USA): 0.8 to 1.2	The impact coefficient has increased over time in the United States and decreased in Canada. The reason could have to do with changes in financial system structure in those countries.
Winker, P. (1999) Germany	Short term: 0.1 (lending rate) and 0.42 (deposit rate) Long term coefficient tends to 1	The speed of adjustment to changes in the money market rate is lower in lending rates than in deposit rate
Manzano and Galmés (1996) Spain	Short term: 0.25–0.75 (lending rate) and 0.2–0.5 (deposit rate) Long term: 0.66–1.2 (lending) and 0.63–0.81 (deposit)	The lending rate tends to response faster in the short and the long run. The type of customer affects the degree of response.

1.2 Chile Compared with Other Countries

This section presents the results at the aggregate level for the Chilean banking industry. The lending rate at the aggregate level was constructed using a weighted average of interest rate for individual banks; the weights were the total amount of loans in the corresponding category. Figure 1 plots the lending interest rate and the interbank rate for the period under analysis. The lending rates follow the interbank interest rate very closely.

An important feature to take into account is that Chilean banks conduct transactions in pesos and in *unidades de fomento* (UF), which is a unit of account indexed to past inflation.[2] This unit of account is used for medium- and long-term transactions. We therefore estimated equation (1) for peso-denominated loans and UF-denominated loans. The most common maturity for the former is less than thirty days (approximately 50 percent of total nominal loans). For the latter, the typical maturity is 90 to 360 days, but it is mainly concentrated around 90 days (approximately 40 percent of total UF-indexed loans). Figure 2 presents the evolution of the lending interest rate for loans of longer maturity and the interest rate on ninety-day Central Bank indexed promissory notes (PRBC). Again, the two interest rates move closely together.[3]

Figure 1. Lending Interest Rate and Interbank Rate

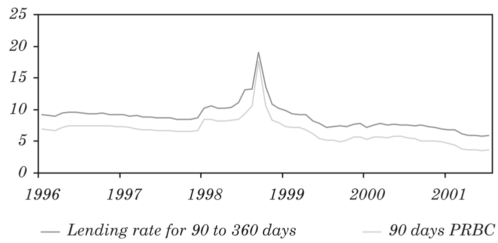


Source: Superintendence of Banks and Financial Institutions (SBIF) and Central Bank of Chile.

2. See Schiller (2002) for a discussion of the use of indexed unit accounts around the world and the UF.

3. Monetary policy is handled through the interbank interest rate, although the ninety-day PRBC interest rate is a good measure of the monetary policy rate for ninety days.

Figure 2. Lending Interest Rate and Ninety-day PRBC



Source: SBIF and Central Bank of Chile.

Next, we estimated a model represented by equation (1). The number of lags chosen was sufficiently high such that the error term becomes white noise. Several papers estimate this equation using different parameterization. The most popular is the error correction model, based on the idea that interest rates are not stationary. There are good economic arguments for disregarding that possibility for interest rates.[4] Nevertheless, to be skeptical, we ran different tests for unit roots, which are presented in the appendix. All the tests reject the presence of unit roots, so we proceeded to run the model in levels.

Table 2 presents the results for the interest rate applied to peso-denominated loans. Columns 1 and 3 show the results of equation (1) controlling for inflation; columns 2 and 4 take into account the dramatic increase in the interest rates during 1998, using a dummy variable, D98, that takes the value one for January 1998 to October 1998. Although the dummy variable is statistically significant, the overall conclusions do not change much. The impact coefficient fluctuates between 0.7 to 0.8, while in all cases the hypothesis of the long-run coefficient being equal to one cannot be rejected. Therefore, on average, banks fully adjust the lending rate to a change in the interbank interest rate in the long run.

Table 3 shows the results for the indexed lending rate. Again, we controlled for the 1998 interest rate turmoil, but it was not statistically significant except for July 1998. The inflation rate was not included, since the variables are indexed interest rates. The impact coefficient is around 0.85, while the long-term coefficient is statistically equal to 1.

4. See Chumacero (2001) for a discussion of unit roots based on economics.

Table 2. Interest Rate Transmission: Nominal Lending Rate[a]

Variable	*30-day lending rate*	*30-day lending rate*	*30- to 89-day lending rate*	*30- to 89-day lending rate*
Interbank rate	0.7932 (14.7964)**	0.8109 (22.8482)**	0.7122 (12.6719)**	0.7098 (18.8454)**
Interbank rate (t – 1)		–0.3355 (–3.8715)**	–0.1670 (–1.8404)	–0.1994 (–2.3729)*
Interbank rate (t – 2)	–0.3129 (–2.3391)*	–0.3193 (–2.9958)**	–0.2659 (–4.4942)**	–0.3330 (–4.1670)**
Interbank rate (t – 3)			0.0750 (2.2498)*	0.0874 (2.3841)*
Interbank rate (t – 4)		–0.0560 (–2.1570)*		
Interbank rate (t – 6)		0.0784 (3.4636)**		
D(MPR)	0.0281 (2.8474)**	0.0259 (3.2080)**	0.0419 (4.0445)**	0.0406 (4.2109)**
Lending rate (t – 1)	0.2865 (3.0554)**	0.5629 (6.1349)**	0.4583 (4.0831)**	0.4059 (4.6310)**
Lending rate (t – 2)	0.2320 (2.2617)*	0.2750 (2.8149)**	0.1896 (2.5192)*	0.3185 (3.2959)**
Inflation (t – 2)	–0.1033 (–2.7302)**	–0.0953 (–3.5682)**	–0.2190 (–4.1982)**	–0.5084 (–3.8040)**
D98				0.4462 (3.9445)**
D98* Interbank rate		–0.3820 (–3.1078)**		
D98* Interbank rate (t – 1)		0.3547 (2.9385)**		
D98*D(MPR)		0.2038 (4.6452)**		–0.1996 (–4.8414)**
Constant	0.1358 (3.8643)**	0.0473 (1.2736)	0.1737 (3.4792)**	0.1538 (3.2508)**
Long-run coefficient (λ)	0.9972	1.1017	1.0060	0.9604
Wald test (λ = 1)	(0.0015)	(0.3202)	(0.0044)	(0.0932)
Summary statistic				
R^2	0.9554	0.9742	0.9466	0.9569

a. t statistics are in parentheses.
* Statistical significance at the 5 percent level.
** Statistical significance at the 1 percent level.

How do these results compare with the international evidence? Table 4 exhibits the comparison between the coefficient reported in column 2 of tables 2 and 3. The estimates for Chile show a high flexibility of the banking interest rate. In fact, the estimation positions Chile close to Mexico and the United Kingdom. According to Cottarelli

Table 3. Interest Rate Transmission: Indexed Lending Rate

Variable	*90- to 360-day lending rate*	*90- to 360-day lending rate*
PRBC	0.8575	0.8553
	(63.3162)**	(48.3335)**
PRBC (t – 1)	–0.4324	–0.2931
	(–4.9115)**	(–4.7812)**
PRBC (t – 2)	–0.0775	–0.0694
	(–5.1854)**	(–3.5892)**
PRBC (t – 4)	0.0357	
	(4.0652)**	
PRBC (t – 5)	–0.0245	–0.1674
	(–1.7402)	(–2.9301)**
Lending rate (t – 1)	0.6396	0.4940
	(6.1577)**	(7.4194)**
Lending rate (t – 5)		0.1643
		(2.8632)**
D98 (July)		1.6035
		(9.1060)**
Constant	0.8019	0.8342
	(3.3145)**	(4.6351)**
Long-run coefficient (λ)	0.9953	0.9520
Wald test (λ = 1)	(0.0757)	(0.0404)
Summary statistic		
R^2	0.9837	0.9924

a. t statistics are in parentheses.
* Statistical significance at the 5 percent level.
** Statistical significance at the 1 percent level.

and Kourelis, the variables that tend to increase the interest rate pass-through are the degree of competition and financial liberalization. It is important to take into account that the time periods are different for the countries included in Cottarelli and Kourelis (1994) with respect to the present study. The former uses data for the 1980s, while we use data for the 1990s. Relevant conditions for interest rate sluggishness were different in the 1990s than in previous decades.

2. Evidence for Chile at the Bank Level

The previous section exposed some evidence in favor of interest rates stickiness. This is the case for almost all the countries that have been studied to date, and it is also the case of Chile, to some

Table 4. International Comparison of Interest Rate Stickiness

Region and country	*Impact*	*Long term*
Latin America		
Chile (nominal rate)	0.81	0.97
Chile (indexed rate)	0.86	0.95
Colombia	0.42	1.03
Mexico	0.83	1.29
Venezuela	0.38	1.48
North America		
Canada	0.76	1.06
United States	0.32	0.97
Europe		
Germany	0.38	1.04
Italy	0.11	1.22
Spain	0.35	1.12
United Kingdom	0.82	1.04

Source: Cottarelli and Kourelis (1994) and authors' estimates for Chile.

extent.[5] We further argued that previous studies suggest that sluggish adjustment is related to market conditions and regulation of the banking sector. In this section, we use bank-level data to explore the factors that may influence the degree of delay in market interest rate response to changes in the policy rate.

For this purpose, we analyze the differences in the interest rate levels charged by banks and the adjustment to changes in the policy rate. In the Chilean case, we observe an important divergence between the interest rates charged by banks, as well as significant differences within a bank depending on the kind of loan, the type of customer, firm or household, and the amount of the loan. Legislation imposes a ceiling on the interest rate charged by loan category, which somewhat limits this dispersion (50 percent above the average market interest rate by loan category).[6]

Our aim is to identify which characteristics might explain the differences in the average rates charged by each bank and their responsiveness to movements in the policy rate. The main characteristics considered were the size of the bank, the type of customer, and the loan risk level. Other variables, such as solvency or liquidity, were also considered, but they did not prove to be significant for explaining differences in lending rates, so the results are not presented.

5. As shown in section 1, the impact effect of changes in the policy rate were less than one for most of the countries studied, including Chile.

6. SBIF (2000).

The data used is at the bank level. We do not have enough information at this point on different transactions within a bank, but this area represents an important future extension of the study.

2.1 Stylized Facts for the Chilean Banking Industry

Tables 5 and 6 show that larger banks charged, on average, lower interest rates than smaller banks during the sample period. For smaller banks, the nominal monthly rate was 1.21, whereas for larger banks this rate was 1.16 for the period 1996–2002. In the case of the UF rate, smaller banks showed a yearly rate of 8.55 percent, on average—that is, 3.5 percent higher than the average for larger banks (8.26 percent). This evidence might support two alternative hypotheses: namely, the structure-performance hypothesis or the efficiency-structure hypothesis. Under the first hypothesis, differences in prices would respond solely to imperfect competition, with differences in price elasticities across markets served by different banks. The second would imply that there are cost advantages for larger banks, together with some degree of market imperfection that allows inefficient banks to survive, at least in the short run.

Table 5. Large Banks: Thirty-day Nominal Rate and Correlation with the Interbank Rate, by Loan Risk and Type of Customer[a]

	Loan Risk[c]		
Type of customer[b] indicator	*< 2 percent*	*> 2 percent*	*Total*
Household loans < 10 percent			
Interest rate			
Correlation			
No. Banks	0	0	0
Household loans > 10 percent			
Interest rate	1.08	1.20	1.16
Correlation	0.90	0.86	0.88
No. Banks	2	4	6
Total			
Interest rate	1.08	1.20	1.16
Correlation	0.90	0.86	0.88
No. Banks	2	4	6

Source: Authors' calculations using data from SBIF.
a. Average for the 1996–2002 period. Large banks are those that have a market share over total loans of more than 5 percent.
b. Type of customer is measured as household loans as a percentage of total loans.
c. Loan risk is measured as past-due loans as a percentage of total loans

Table 6. Small Banks: Thirty-day Nominal Rate and Correlation with the Interbank Rate, by Loan Risk and Type of Customer[a]

	Loan Risk[c]		
Type of customer[b] indicator	*< 2 percent*	*> 2 percent*	*Total*
Household loans < 10 percent			
Interest rate	1.12	1.37	1.19
Correlation	0.83	0.76	0.81
No. Banks	5	3	8
Household loans > 10 percent			
Interest rate	1.25	1.21	1.23
Correlation	0.87	0.79	0.83
No. Banks	3	3	6
Total			
Interest rate	1.17	1.27	1.21
Correlation	0.85	0.78	0.82
No. Banks	8	6	14

Source: Authors' calculations using data from SBIF.
a. Average for the 1996–2002 period. Small banks are those that have a market share over total loans of less than 5 percent.
b. Type of customer is measured as household loans as a percentage of total loans.
c. Loan risk is measured as past-due loans as a percentage of total loans

In terms of loan risk, banks with a higher percentage of past-due loans (more than 2 percent) charged, on average, higher interest rates to their clients, as expected. This is 11.1 percent higher in the case of nominal rates and 8.6 percent in the case of UF rates, over the sample period. When we compute a simple correlation between lending rates and our indicator for the policy rate (the interbank rate in the case of nominal interest rates and the ninety-day PRBC in the case of UF interest rates), this correlation is smaller for banks with lower-quality loans. This may be due to adverse selection problems, in the sense that if interest rates increase, only riskier projects (with a higher expected return) would stay in the market and the average quality of the loan portfolio would decrease, thereby lowering the bank's profits. Banks will thus not respond rapidly to an increase in the policy rate, especially in the case of banks with a higher portion of past-due loans. On the other hand, if the policy rate decreases, we would expect less responsiveness from banks with a riskier portfolio, because it is more difficult for riskier clients to move to other banks. Banks with a larger portion of past due loans thus have less incentive to decrease interest rates, at least in the short run.

Finally, in tables 7 and 8 we analyze differences in interest rates charged by banks classified by type of loan.[7] We are able to make this distinction only for smaller banks because larger banks do not display much difference within this category, since all of them have more than 10 percent of household loans. So, for smaller banks we have two groups: those with less than 10 percent of the loans given to households and those with more than 10 percent.

In the case of both nominal interest rates and UF interest rates for smaller banks, the higher average rate charged corresponds to banks that have a larger portion of past-due loans and a lower share of household loans, while banks with low risk and a low share of household loans charge lower interest rates. This indicates that there is an important dispersion of interest rates charged to firms, which seems to be larger than in the case of households. This evidence suggests that the demand elasticity of households is larger than that of firms. A possible explanation for this is that asymmetric information leads firms to establish a long-term relationship with their banks to a greater extent than households; this gives additional market power to the banks, owing to higher switching costs for firms.

Table 7. Large Banks: Ninety-day to One-year Indexed Rate and Correlation with the PRBC Rate, by Loan Risk and Type of Customer[a]

	Loan Risk[c]		
Type of customer[b] indicator	*< 2 percent*	*> 2 percent*	*Total*
Household loans < 10 percent			
Interest rate			
Correlation			
No. Banks	0	0	0
Household loans > 10 percent			
Interest rate	8.02	8.38	8.26
Correlation	0.95	0.94	0.95
No. Banks	2	4	6
Total			
Interest rate	8.02	8.38	8.26
Correlation	0.95	0.94	0.95
No. Banks	2	4	6

Source: Authors' calculations using data from SBIF.
a. Average for the 1996–2002 period. Large banks are those that have a market share over total loans of more than 5 percent.
b. Type of customer is measured as household loans as a percentage of total loans.
c. Loan risk is measured as past-due loans as a percentage of total loans

7. The type of loan is measured as the percentage of total loans made to households (consumption plus mortgage).

Table 8. Small Banks: Ninety-day to One-year Indexed Rate and Correlation with the PRBC Rate, by Loan Risk and Type of Customer[a]

	Loan Risk[c]		
Type of customer[b] indicator	*< 2 percent*	*> 2 percent*	*Total*
Household loans < 10 percent			
Interest rate	8.17	9.14	8.52
Correlation	0.92	0.80	0.87
No. Banks	5	3	8
Household loans > 10 percent			
Interest rate	8.38	8.80	8.59
Correlation	0.91	0.94	0.92
No. Banks	3	3	6
Total			
Interest rate	8.25	8.96	8.55
Correlation	0.92	0.87	0.90
No. Banks	8	6	14

Source: Authors' calculations using data from SBIF.
a. Average for the 1996–2002 period. Small banks are those that have a market share over total loans of less than 5 percent.
b. Type of customer is measured as household loans as a percentage of total loans.
c. Loan risk is measured as past-due loans as a percentage of total loans.

2.2 A Model for Lending Rate Stickiness

This section presents a model that we use to build on some of the hypotheses that we test for the Chilean banking industry. These hypotheses are related to the stylized facts presented in the previous section. The model gives us some insights about what to expect from our empirical analysis, as well as possible explanations for our findings.

It seems appropriate to assume an imperfect competition model in the case of the banking sector, where there are significant barriers to entry and an important degree of product differentiation.[8] We also assume that there is asymmetric information in this industry, which leads to adverse selection and moral hazard problems. We combine these two issues by assuming that banks make a two-step decision, which considers the long-run equilibrium and the short-run behavior that will take them to this condition.[9]

For the long run, we assume a simple Monte-Klein model for a monopolistic bank that faces a downward sloping demand for loans

8. Freixas and Rochet (1998).
9. This method of combining these two factors is similar to Scholnick (1991), Winker (1999), and Bondt (2002).

$L(i_L)$ and an upward sloping supply of deposits $D(i_D)$. This captures the fact that banks have some monopoly power. The decision variables for the firm are the quantities of loans (L) and deposits (D). Bank k maximizes the following profit function:

$$\pi_k(L,D) = \left[\gamma_k i_{L,k}(L) - m\right]L_k + \left[m(1-\alpha) - i_{D,k}(D)\right]D_k - C(D_k, L_k)\ , \qquad (3)$$

where γ_k is the probability that the loan will be repaid, m is the interbank rate (which is given for individual banks), α is the proportion of deposits that constitutes cash reserves, i_D is the deposit interest rate, and i_L is the lending interest rate. $C(D,L)$ accounts for the total cost of intermediation services, which is a function of the total amount of deposits and loans.

Solving for the first-order conditions and rearranging terms, we get to the following expressions for the lending interest rate:

$$i_L^* = \frac{\varepsilon_k}{(\varepsilon_k - 1)\gamma_k}\left(m + C_L'\right)\ , \qquad (4)$$

where ε_k is the absolute value of the demand elasticity for loans, which is greater than 1 since we are assuming monopolistic competition. For the purpose of this paper, we are interested in the loan market and we assume that costs are separable, so that the optimal lending rate is independent of the characteristics of the deposit market. This simple model leads us to conclude that different interest rates charged on loans may reflect different demand elasticities and the probability of loan repayment (portfolio risk).

The above model is interpreted as the long-run equilibrium for banks. To simplify our model, we assume each bank faces a constant elasticity demand function. In other words, ε might be different for each bank, but it is independent of i_L. We can write this relationship between the lending rate and the interbank rate as $i_L^* = \Phi_k m$. (Here, $\Phi_k = \varepsilon_k / (\varepsilon_k - 1)\gamma_k$ is a mark up, which is a function of demand elasticity and the repayment probability). Thus, the long-run pass-through coefficient is larger the smaller is the demand elasticity and the smaller is the probability of repayment. This long-run coefficient may or may not be equal to 1, when there is some degree of monopoly power.

Asymmetric information, however, results in a sluggish adjustment process to get to this long-run equilibrium. In fact, we are interested in finding out whether there is some delay in the response of market interest rates to changes in the policy rate and whether this delay depends on bank characteristics related to demand elasticity and asymmetric information.

Specifically, we are thinking of a setup in which in the short run, banks solve an intertemporal problem characterized by a cost of adjusting too slowly to this long-run equilibrium and a cost of moving too fast. This latter cost is due to adverse selection and moral hazard problems in the banking industry. For instance, if a bank increases the lending rate in response to an increase in the money market rate, the bank's adjustment to its new long-term equilibrium may involve attracting debtors that have a lower repayment probability, thereby lowering the bank's profits. At the same time, moral hazard arises because a higher interest rate gives debtors incentives to invest in riskier projects, which would also decrease the bank's profits.[10] Under this framework, therefore, we assume that there are some adjustment costs stemming from asymmetric information. This is modeled as a quadratic loss function following Nickell (1985), Scholnick (1991), and Winker (1999), which is tractable because it generates a linear decision rule.[11] The loss function for bank k in period t is the following:

$$\Gamma_{t,k} = \sum_{s=0}^{\infty} \delta^s \left[\omega_{1,k} \left(i_{k,L,t+s} - \Phi_k m_{t+s} \right)^2 + \omega_{2,k} \left(i_{k,L,t+s} - i_{k,L,t+s-1} \right)^2 \right], \tag{5}$$

where ω_1 and ω_2 represent the weight that the bank gives to achieving the long-run target value for the lending rate and the cost of moving to that target value, respectively. Recall that Φ_k is a function of the demand elasticity and the probability of repayment that bank k faces, whereas ω_j, j = 1,2, depends on the bank's average loan risk. If the portion of past-due loans for bank k is higher, the adverse selection or moral hazard problem for that bank becomes more important and the bank will give more weight to changes in the interest rate, which implies a slower adjustment. On minimizing equation (5), we obtain

$$i_{k,L,t+s} = \frac{\omega_{1,k}}{\omega_{1,k} + \omega_{2,k}} \Phi_k m_{t+s} + \frac{\omega_{2,k}}{\omega_{1,k} + \omega_{2,k}} i_{k,L,t+s-1}. \tag{6}$$

10. Stiglitz and Weiss (1981).

11. Scholnick (1991) and Winker (1999) also include a third term in the loss function, but it is not included in our setup. For an argument, see Nickell (1985). The other difference is that we have a multiplicative mark-up instead of an additive mark-up.

Equation (6) shows that the impact coefficient depends on the size of $\omega_{1,k}$ relative to $\omega_{1,k} + \omega_{2,k}$ and the mark up, Φ_k. Therefore, the long-run coefficient is always larger than the short-term coefficient. The bank's loan risk determines Φ_k and $\omega_{2,k}$: the lower the probability of repayment (higher risk), the higher are both Φ_k and $\omega_{2,k}$. If the debtors are too risky and the effect on $\omega_{2,k}$ is more important, the bank may not completely pass through a money market interest rate increase (in the short run) because it would stifle the debtors. In the long run, however, the interest rate charged will reflect the risk characteristic of the debtor. In other words, unpaid loans should have a negative effect on the impact coefficient and a positive effect on the long-term multiplier.

The main difference between our setup and the one presented by Scholnick (1991) and Winker (1999) is that they derive an error correction model (ECM) from this quadratic loss function. Our variables are stationary, however, even if we assume that there is a long-run relationship between the interbank rate and the lending rate. We therefore estimate our econometric model in levels and not in an ECM form. Recall that the ECM has this interpretation only if the variables are nonstationary and cointegrated, which is not the case for our data.[12]

The other important difference is that we use the above model in a panel data estimation (in section 2.3) that allows the parameters to be different for different banks depending on their characteristics.

2.3 Econometric Results

The model described above suggests that differences in interest rate pass-through might be related to product characteristics such as the type of customer or the risk level of the loan portfolio. The econometric analysis presented in this section allows us to address this issue by estimating a dynamic panel data model in which bank characteristics are interacted with the interbank rate and its lags. An alternative method is time series estimation by bank, but it has the drawback that changes in bank characteristics during this time may be affecting the sluggishness of adjustment for each bank, which is not correctly captured.[13]

12. Unit root tests is presented in the appendix. Derivation of the ECM and explanation of why it is not appropriate with stationary data are found in Nickell (1985) and Wickens and Breush (1988).

13. See Berstein and Fuentes (2003) for time series estimations at the bank level.

We estimate the following equation, which is based on the model described in section 2.2. Adverse selection is captured by the adjustment cost coefficient of the model, which is a function of the quality of loan portfolio, and we allow demand elasticity to be a function of the type of customers the bank has and the size of the bank.

$$i_{h,t} = \eta_h + \sum_{j=1}^{m} \beta_j i_{h,t-j} + \sum_{k=0}^{n} \theta_k m_{t-k} + \sum_{k=0}^{n} \alpha_k X'_{h,t-k} M_{t-k} + \sum_{l=0}^{p} \phi_{t-l} \Delta MPR_{t-l} \qquad (7)$$

where M is a vector that contains the lagged lending rate and the money market rate, and X' is a vector of bank characteristics, which includes: the loan portfolio risk, measured as the portion of past-due loans; type of consumers, measured as the share of household loans (consumption and mortgage); and bank size, measured as the percentage of total loans. Finally, η_h is a bank-specific effect.

The problem of estimating dynamic panel data has been widely discussed in the literature, and different methods have been proposed to obtain consistent estimates of the parameters. Anderson and Hsiao (1982) propose a method based on instrumental variables (IV), which consists in taking first differences of the equation to eliminate unobserved heterogeneity and then using instrumental variables to estimate consistently the parameters of the lag-dependent variables.

For instance, the following equation is to be estimated using panel data:

$$y_{it} = \rho y_{it-1} + \beta x_{it} + \eta_i + u_{it} \ , \qquad (8)$$

where y_{it} represents the lending interest rate, x_{it} represents a dependent variable like the interbank interest rate, and η_i is the unobserved heterogeneity. After taking the first difference, the equation to be estimated is

$$y_{it} - y_{it-1} = \rho(y_{it-1} - y_{it-2}) + \beta(x_{it} - x_{it-1}) + u_{it} - u_{it-1} \ . \qquad (9)$$

Anderson and Hsiao propose $y_{i,t-2}$ or $(y_{i,t-2} - y_{i,t-3})$ as an instrument for $(y_{i,t-1} - y_{i,t-2})$, but Arellano (1989) shows that $y_{i,t-2}$ is a much better instrument for a significant range of values of the true in equation (9).

Arellano and Bond (1991) propose an alternative methodology based on generalized method of moments (GMM) estimators. This method uses several lags of the variables included as instruments, so it is especially efficient when T is small and N is large.[14] The method is applied to equation (6), using moment restrictions that come from the use of instrumental variables. Judson and Owen (1999) provide evidence that for small T, GMM is a better estimator than Anderson and Hsiao's methods under the mean square error criterion. It is unclear, however, which method is better for unbalanced panel data and T around 20.

Several other methods have been developed based on the traditional within-group, IV, and GMM estimators. The IV method tends to work better than the within-group estimator when N tends to infinity (N is very large) and T is fixed. Alvarez and Arellano (2002) show the asymptotic property of the within-group, GMM, and limited information maximum likelihood (LIML) estimators. An important result for our case is that regardless the asymptotic behavior of N, the estimator of ρ is consistent when T goes to infinity. Moreover, if $\lim(N / T) = 0$ (as T goes to infinity) there is no asymptotic bias in the asymptotic distribution of the within-group estimator, while in the opposite case of $\lim(T / N) = 0$ (as N goes to infinity), there is no asymptotic bias in the asymptotic distribution of the GMM estimator. In our panel, T is large and increasing over time, while N remains relatively fixed. The traditional within-group estimator will thus provide the best results.[15]

Tables 9 and 10 show the results for the thirty-day nominal interest rate and for the 90- to 360-day indexed interest rates, respectively. The first column of tables 9 and 10 present the results of the panel estimation without controlling for the 1998 effect and without considering the interaction between bank characteristics and the right-hand-side variables. If we compare these regressions with the ones from section 1, we observe that the impact and long-run effects (shown at the bottom of each table) are smaller than what we found previously. Note that previously, we were estimating impact and long-run effects at an aggregate level using the weighted average interest rates, so that large banks drive the results to a larger extent on those regressions than on the panel data estimation.

14. See Judson and Owen (1999) for further discussion of the advantages of different methodologies.

15. See Berstein and Fuentes (2003) for panel data estimations using Anderson and Hsiao, and Arellano and Bond methods.

Table 9. Panel with Interaction and 1998 Dummies, Thirty-day Nominal Rate

Explanatory variable	*(1)*	*(2)*	*(3)*
Interbank rate	0.74	0.72	0.74
	(41.51)**	(34.80)**	(24.92)**
Interbank rate (–1)	–0.30	–0.41	–0.48
	(–10.86)**	(–14.44)**	(–13.02)**
Interbank rate (–5)	–0.12	–0.06	
	(–6.91)**	(–3.84)**	
Interbank rate (–6)	–0.06		
	(–2.43)**		
Nominal rate, 30 days (–1)	0.57	0.67	0.68
	(26.84)**	(32.80)**	(28.36)**
Nominal rate, 30 days (–3)	0.05		
	(3.44)**		
Nominal rate, 30 days (–6)	0.14	0.06	0.04
	(6.58)**	(4.05)**	(2.72)**
D (MPR)	0.04	0.03	0.06
	(8.62)**	(5.71)**	(7.08)**
Inflation			
Inflation (–2)	–0.13	–0.08	–0.09
	(–6.83)**	(–4.60)**	(–3.65)**
Interbank * risk (–1)			–2.31
			(–2.13)*
Interbank (–1) * risk (–2)			5.05
			(4.80)**
Interbank (–1) * market share (–1)			–0.72
			(–2.84)**
Interbank * Cons.			0.18
			(1.77)
Long-run coefficient	1.07	0.88	1.09
Standard deviation	(0.07)	(0.06)	(0.08)
Summary statistic			
No. observations	1,447	1,447	1,105
No. banks	20	20	20

a. The dependent variable is the thirty-day nominal interest rate. Models (2) and (3) control for the year 1998. The models were estimated using fixed effects, which are not reported; *t* statistics in parentheses.
* Statistical significance at the 5 percent level.
** Statistical significance at the 1 percent level.

The second column of tables 9 and 10 present the results of the panel estimation controlling for the 1998 effect. The impact and the long-run coefficients decrease relative to those reported in the first column of each table, but the values are consistent with the idea that the long-term coefficient is larger than the short-term coefficient. However the long-term coefficient is not statistically equal to 1. The

Table 10. Panel with Interaction and 1998 Dummies, Ninety-day to One-year Indexed Rate[a]

Explanatory variable	*(1)*	*(2)*	*(3)*
PRBC	0.88	0.71	0.72
	(90.95)**	(37.19)**	(25.41)**
PRBC (–2)	0.05	–0.03	
	(2.62)**	(–2.30)*	
PRBC (–3)	–0.38	–0.21	–0.21
	(–12.22)**	(–7.61)**	(–6.54)*
PRBC (–4)	–0.09		
	(–3.10)**		
PRBC (–5)	–0.05	–0.13	–0.13
	(–3.98)**	(–4.98)**	(–4.18)**
PRBC (–6)	–0.09	–0.05	
	(–3.31)**	(–1.96)*	
	0.25	0.24	0.19
UF rate, 90 days to 1 year (–1)	(14.10)**	(19.36)**	(12.13)**
	0.26	0.24	0.24
UF rate, 90 days to 1 year (–3)	(9.41)**	(9.52)**	(8.02)**
	0.09		
UF rate, 90 days to 1 year (–4)	(3.19)**		
UF rate, 90 days to 1 year (–5)		0.12	0.12
		(4.32)**	(4.20)*
UF rate, 90 days to 1 year (–6)	0.09	0.05	
	(3.47)**	(2.19)*	
D [MPR (–1)]	–0.34		
	(–6.14)**		
PRBC (–2) * risk (–3)			–2.48
			(–4.12)**
UF rate, 90 days to 1 year (–1) *			1.47
risk (–2)			
			(3.34)**
PRBC * market share			–0.34
			(–3.11)**
PRBC (–2) * Cons.(–2)			0.18
			(3.83)**
Long-run coefficient	1.04	0.84	0.85
Standard deviation	(0.03)	(0.03)	(0.04)
Summary statistic			
No. observations	1,368	1,368	1,368
No. banks	18	18	18

a. The dependent variable is the thirty-day nominal interest rate. Models (2) and (3) control for the year 1998. The models were estimated using fixed effects, which are not reported; *t* statistics in parentheses.
* Statistical significance at the 5 percent level.
** Statistical significance at the 1 percent level.

last column in each table allows us to check the hypotheses provided by the theoretical model. In the case of nominal interest rates, the riskier the portfolio, the lower is the impact coefficient, which is consistent with the idea that banks will not pass interest rate change on to debtors in the short run, according to the difference equation (6). In the long run, however, the pass-through will be larger the riskier is the portfolio. This relationship is represented in figures 3 and 4, which illustrate how the average loan risk has increased over time and the estimated impact effect has decreased while the long run effect gets larger.

In the case of the indexed interest rate, the results are different. The impact coefficient is not affected by the portfolio risk, while the level of the unpaid loans affects the long-run coefficient by reducing it (see figure 5).

Figure 3. Impact Effect and Loan Risk, Thirty-day Nominal Rate

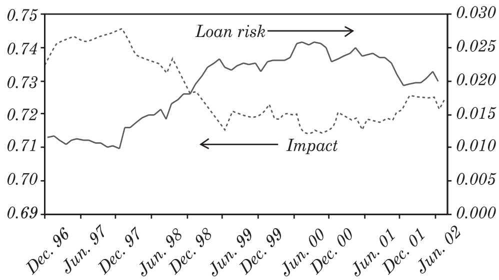


Figure 4. Long-run Effect and Loan Risk, Thirty-day Nominal Rate

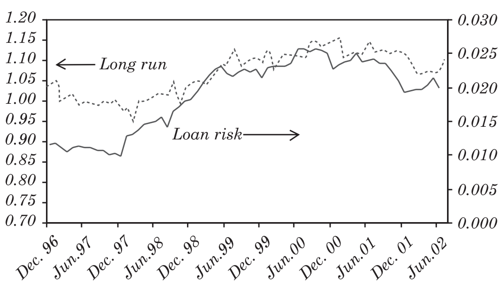
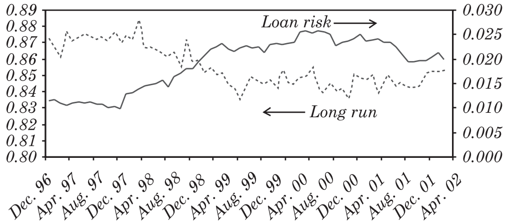


Figure 5. Long-run Effect and Loan Risk, Ninety-day to One-year UF Rate



Finally, for both nominal and indexed rate, bank size negatively affects the pass-through, while banks that are more oriented toward households have a larger pass-through.

3. Concluding Remarks

The estimates presented in this paper support the fact that the banking interest rate in Chile is highly flexible. In fact, the estimation positions Chile close to Mexico and the United Kingdom, countries displaying the highest degree of flexibility.

An earlier study by Cottarelli and Kourelis (1994) identifies the degree of competition and financial liberalization as the main determinants of interest rate stickiness. We used bank-level data to explore other factors that influence the degree of delay in market interest rate response to changes in the policy rate. The main characteristics identified in our analysis of differences in the interest rate levels charged by banks and their adjustment to changes in the policy rate are bank size, type of customers, and the loan risk level.

Our bank-level econometric analysis found significant differences in banks' responses to changes in the policy interest rate. Moreover, the smaller the bank, the lower the portion of past-due loans, and the larger the share of household consumers—the faster is the response of lending interest rates to movements in the money market rate. These results are consistent with the model and the stylized facts presented in the paper.

Topics for future research include alternative measures for capturing loan risk and other characteristics that would help improve measures of different demand elasticities, at the bank level. Furthermore, the availability of disaggregated information on the interest rates charged for different types of loans within a bank would improve estimates of the effects of loan risk or type of customer on the interest rate responses to changes in policy rates.

Appendix
Unit Root Tests

We ran different tests for unit roots, all of which reject the presence of unit roots. The results are presented in table A1. The tests consider a trend for the nominal interest rates, and we used the modified Akaike information criterion to choose the number of lags. We use augmented Dickey-Fuller (ADF) and Phillips-Perron tests with the modified Akaike to solve the size problem of the tests, but the power is very low. The power of the tests is higher when using Dickey-Fuller generalized least squares (DF-GLS) and Phillips-Perron-Ng.

Table A1. Unit Root Tests, 1995 to 2001

Rate	*ADF*	*DF-GLS*	*Phillips-Perron*	*Phillips-Perron Ng Mzt*
PRBC	–1.928	–1.949*	–2.630	–1.995*
Interbank rate	–3.733*	–3.175*	–4.364**	–3.135*
UF, 90 days to 1 year	–2.258	–2.292*	–2.204	–2.134*
Nominal rate, 30 days	–4.169***	–4.612**	–4.686**	–3.562**

* Nonstationarity rejected at 5 percent.
** Nonstationarity rejected at 1 percent.

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