

A Theoretical and Empirical Assessment of the Bank Lending Channel and Loan Market Disequilibrium in Poland

Christophe Hurlin*, Rafał Kierzenkowski**¹

Abstract

We study the impact of the bank lending channel and loan market disequilibrium on the efficiency of the monetary policy transmission in Poland since 1994. First, we develop a simple credit-augmented model with an interest rate control, flexible prices and an imperfect nominal wage indexation. Within this framework, we establish that the bank lending channel may amplify but also attenuate the impact of monetary policy shocks on output and prices as compared to the traditional interest rate channel. The variations in the interest rate spread between the loan rate and the central bank's intervention rate are a good indicator when distinguishing between amplification and attenuation effects of monetary policy shocks provided that there is a positive relationship between both rates and that the loan interest rate is a market clearing variable. Second, we apply a regime switching framework to the loan market. The results suggest that disequilibrium is a permanent characteristic of the Polish loan market since 1994. Moreover, we discuss empirically the impact of any type of disequilibrium in the loan market on the effectiveness of the bank lending channel. We find attenuation effects of the bank lending channel on monetary policy shocks from the beginning of 1996 to August 1998, and on average a neutral effect of this transmission channel from September 1998 to June 2001.

¹ This is a revised version of a paper presented at Warsaw School of Economics (Chair of Monetary Policy) on December 4, 2001 and at the National Bank of Poland (Research Department) on December 11, 2001. The authors are grateful to Michał Brzoza-Brzezina, Tomasz Chmielewski, Maciej Dudek, Tomasz Łyziak, Bogusław Pietrzak, Zbigniew Polański, Jerzy Pruski, Ewa Wróbel, Marzena Zaremba, and many other participants of both seminars for numerous valuable comments. We would also like to thank Jérôme de Boyer des Roches, Balázs Égert, Kate Finn, Olivier Grosse, Hélène Lenoble-Liaud, Joël Métais, Jean-Marie Renaud and Jérôme Sgard for helpful suggestions. Finally, many thanks go to the following people who provided an inestimable help in obtaining the time series used in this paper: Jakub Borowski, Michał Brzoza-Brzezina, Norbert Cieśla, Marta Gołajewska, Paulina Krzysztofik, Małgorzata Pawłowska, Zbigniew Polański and Paweł Wyczański. The standard disclaimer applies. Comments are welcome.

* EURISCO, Paris IX Dauphine University, and CEPREMAP. E-mail: christophe.hurlin@dauphine.fr

** CREFED-CERPEM, Paris IX Dauphine University. E-mail: rafal.kierzenkowski@dauphine.fr

Place du Marchal De Lattre de Tassigny 75775 Paris.

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Introduction

The transmission mechanism describes the link between monetary policy actions and their impact on real economic activity and inflation. Of course, several interrelated transmission channels may be at work. Yet, it is widely accepted that the Polish financial system is principally a bank-oriented one. This motivates our study since we seek to explain the role the banking sector plays in the transmission mechanism in Poland since 1994. More specifically, we investigate the importance of the bank lending channel and evaluate the disequilibrium in the Polish loan market. The difficulties of the authorities' control over credit activity prove that the Polish banking sector is a key element in understanding the efficiency of monetary policy actions during the 1990s (Polański, 1998; Brzoza-Brzezina, 2000).

Following Bernanke and Blinder's (1988a,b) seminal article, the main result presented in the bank lending channel literature states that the imperfect substitutability between bonds and loans generates an amplification of monetary policy shocks when compared to the traditional money (or interest rate) channel. The bank lending channel makes monetary policy more restrictive (expansionary) than in a standard IS/LM model because of an independent effect that emanates from the asset side of the banking sector, which reduces (increases) the loan supply to "bank-dependent" borrowers². The variations in both the credit supply and the spread between loan and bond interest rates summarize the amplifying nature of the bank lending channel: the interest rate spread increases (decreases) and the supply of credit decreases (increases) in the event of a restrictive (expansionary) monetary policy (Bernanke, 1993).

Kierzenkowski (2001) makes a critical assessment of Bernanke and Blinder's results, demonstrating that they are not general since they require special assumptions (see Bernanke and Blinder (1988a) for their detailed exposition). The bank lending channel can either amplify or attenuate the effects of the traditional interest rate channel. He establishes that, as a general rule, the direction of change in the spread between loan and bond interest rates after a monetary policy shock is a good indicator for distinguishing between these two effects. Following a monetary tightening (expansion) there is an increase (decrease) in the interest rate spread in the event of amplification effects and a decrease (increase) when monetary policy shocks are attenuated.

However, these testable implications cannot be used for empirical investigations in Poland, since Polish monetary authorities use an interest rate and not, as assumed in the model, a base money target policy. Therefore, in section 1, we develop a simple aggregate-demand-and-supply (hereafter AD/AS) credit-augmented model more in line with the conduct of the monetary policy in Poland, assuming an interest rate monetary control, flexible prices of goods and an imperfect nominal wage indexation. In section 2, we apply the testable implications of the model to provide an assessment of the bank lending

² See, for instance, Kashyap, Stein and Wilcox (1993).

channel in Poland.

An empirical identification of a disequilibrium in the loan market is of primary importance for the conduct of monetary policy. The disequilibrium results from market imperfections leading to an incomplete price adjustment of the loan interest rate and therefore to a possible distortion of monetary policy impulses. We deal with this issue estimating a regime-switching model that allows for two regimes in order to characterize the annual growth rate of the quantity of loans extended to Polish firms. A demand (supply) regime occurs if the growth rate of the quantity of loans is determined by the variables and their parameters associated with the annual increase in loan demand (supply). In section 3, we precisely describe the theoretical methodology used in the paper, outlining different points that one must be aware of in order to get consistent estimators. In section 4, we present the specification research and the final results that we analyze in the Polish monetary policy context.

In our theoretical model of transmission we assume that the loan interest rate is perfectly flexible, thus clearing the loan market. This is a standard assumption made in the bank lending channel literature³. Therefore, in section 5, we investigate empirically whether the existence of a loan market disequilibrium precludes the action of the aforementioned transmission channel.

1 A Simple Model of the Bank Lending Channel

1.1 General Assumptions

In the Bernanke and Blinder's (1988a,b) model, monetary policy is characterized in terms of the authorities' control over banking reserves, assuming fixed prices. We extend this framework in several ways.

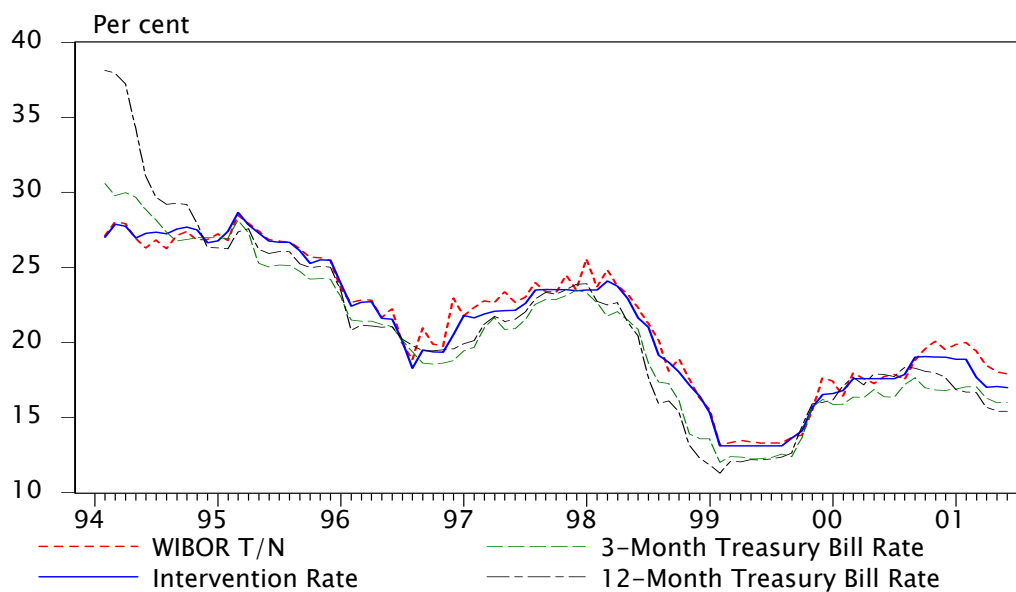
First, considering a perfectly deterministic environment without any stochastic disturbances we invert the policy rule, modelling the central bank as operating on interest rates rather than controlling the base money. The interest rate control assumption reflects the actual conduct of monetary policy in Poland since 1994. According to Osiński (1995, 1999) and Sławiński and Osiński (1997,1998), the National Bank of Poland (hereafter NBP) was setting a 1-day reverse repo interest rate (and more generally was controlling the short-term WIBOR T/N⁴ interest rate) in the 1994-1995 period, while during the 1996-1997 period the main interest rate instrument was a 14-days reverse repo rate. Since February 1998, the basic instrument set by the Monetary Policy Council is represented by the minimum yield on 28-days NBP bills. For the period under consideration (February 1994 - June 2001), these interest rates were used in open-market operations in order to mop up the excess liquidity of the

³ See, for example, Bernanke and Blinder (1988a,b), Kashyap, Stein and Wilcox (1993), Gambacorta (1998).

⁴ Tomorrow Next Warsaw Interbank Offer Rate

banking system created by a combination of strong capital inflows and fixed exchange rate policies followed till late 1990s. We calculated a single intervention rate as a weighted average of 1 to 14-days reverse repo operation rates and that of the central bank securities issued for different maturities between February 1994 and January 1998 and, since then equal to the actual rate on 28-days NBP bills⁵. As it appears in Figure 1, our indicator of monetary policy stance is almost equal to WIBOR T/N and, since at least August 1994, is very close to the yield of 3-month and 1-year Treasury bills on the primary market.

Figure 1. Intervention, WIBOR T/N and Treasury Bills Interest Rates, II/94 - VI/01



Source: National Bank of Poland and the authors' calculations.

An indicator of monetary policy stance comparable to ours is used by Kokoszcyński (1999). Moreover, similarly to Kokoszcyński (1999), we find a significant impact of our indicator on Treasury bills interest rates. More precisely, as shown in Table 1, the intervention rate (*IC*) Granger caused the 3-month Treasury bill interest rate (*IB3M*) for the entire period under consideration, the 6-month interest rate (*IB6M*) in the February 1994 - August 1998 period⁶ but failed to affect the 12-month rate (*IB12M*). However, in the latter case, the expected relationship still occurred for a shorter period of time. On the whole, by controlling its intervention rate, the central bank exerts an important influence on the market interest rates. Given these different observations, we assume, for the sake of simplicity, that the bond interest rate of the model⁷ is equal to the yield of NBP's securities, i.e. to the intervention

⁵ The indicator also includes the average rate of outright operations, which were systematically used since September 2000 and seldom before that date.

⁶ Due to breaks in data since August 1998, the test could not be made for the entire period.

⁷ Empirically, Bernanke and Blinder (1988a,b) use the 3-month Treasury bill interest rate as a proxy for the bond interest rate.

rate. Presenting the model, we use both terms interchangeably.

Table 1. Granger Causality Tests

Null Hypothesis	F-Statistic	Probability	Period
<i>IB3M</i> does not Granger Cause <i>IC</i>	2.478	0.119	II/1994 - VI/2001
<i>IC</i> does not Granger Cause <i>IB3M</i>	4.345	0.040	
<i>IB6M</i> does not Granger Cause <i>IC</i>	0.887	0.350	II/1994 - VIII/1998
<i>IC</i> does not Granger Cause <i>IB6M</i>	5.035	0.029	
<i>IB12M</i> does not Granger Cause <i>IC</i>	3.554	0.062	II/1994 - VI/2001
<i>IC</i> does not Granger Cause <i>IB12M</i>	0.392	0.532	
<i>IB12M</i> does not Granger Cause <i>IC</i>	1.501	0.226	I/1995 - XII/1998
<i>IC</i> does not Granger Cause <i>IB12M</i>	3.860	0.055	

As in Kokoszcyński (1999), we used one lag nominal variables in first differences.

Second, we assume that the prices of goods are perfectly flexible but there is an imperfect indexation of nominal wages to the price level. As a consequence, a monetary policy shock will act on both output and prices. It should be noted, however, that if the central bank is setting, as we assume, the nominal interest rate, this creates a price level indeterminacy problem if prices of goods and nominal wages are both perfectly flexible⁸.

Third, as is standard in the literature, we introduce in the AD/AS framework a bank lending channel working over and above the interest rate channel by assuming that bonds and loans are imperfect substitutes. Therefore, there is a clear distinction between both assets. Hence, following a monetary tightening, banks cannot offset a decline in deposits by simply adjusting their bond holdings and keeping their loan supply unaffected. Similarly, firms cannot offset a decrease in loan supply by simply increasing their bond issue without incurring higher costs.

Finally, as the methodology employed in the model is comparative statics, the expected inflation rate is assumed fixed and omitted.

The characteristics of different markets are as follows.

The loan supply is deduced from the following simplified banks' balance sheet (which ignores net worth):

$$R^b + B^b + L^s = D^s,$$

with assets: nominal reserves, R^b ; nominal bonds, B^b ; nominal loans, L^s ; and liabilities: nominal deposits, D^s . Since reserves consist only of required reserves, i.e. $R^b = \tau D^s$, where τ denotes the reserve requirement coefficient, the banks' adding-up constraint is:

$$B^b + L^s = (1 - \tau)D^s.$$

Assuming that the desired structure of banks' portfolio is a function of rates of return on loans and

⁸ The result of price level indeterminacy of the nominal interest rate instrument in a closed-economy framework under rational expectations was first derived by Sargent and Wallace (1975).

bonds, the loan supply is:

$$L^s = \Gamma(I_l, I_b)D^s(1 - \tau) \quad \text{with: } \Gamma_{I_l} > 0, \Gamma_{I_b} < 0, \quad (1)$$

where Γ is the proportion of deposits out of required reserves that banks wish to hold under credit form. The loan supply is an increasing function of the loan interest rate. This means that the price of loans is perfectly flexible and clears the loan market. Due to the substitution effect, it is a decreasing function of the bond interest rate. In order to simplify our expressions we write hereafter each variable as a deviation around the steady state: we write, for instance for an X variable, x as a deviation in percentage (or in logarithm):

$$x = \log \frac{X}{X_0} \simeq \frac{X - X_0}{X_0}.$$

Therefore, for a given reserve requirement coefficient, the linear form of the loan supply function (1) is:

$$l^s = \gamma_l i_l - \gamma_b i_b + d^s, \quad (2)$$

with γ_l and γ_b denoting the loan interest rate and the bond interest rate elasticities of loan supply respectively. In the credit market, borrowers choose between loans and bonds according to the interest rates on the two instruments. The nominal loan demand is:

$$l^d = p - \lambda_l i_l + \lambda_b i_b + \lambda_y y, \quad (3)$$

with λ_l , λ_b and λ_y standing for the loan interest rate, the bond interest rate and the income elasticities of loan demand respectively, y the real output and p the price of output. The positive dependance on income captures the transactions demand for credit, which might arise from working capital or liquidity considerations.

We ignore cash and we do not model the deposit supply while assuming that it is determined by shocks to deposit demand. Hence, the nominal supply of deposits is equal, for a given reserve requirement ratio, to bank reserves r_b :

$$d^s = r_b. \quad (4)$$

The nominal demand for deposits d^d depends positively on the real income and negatively on the bond interest rate:

$$d^d = p + \beta_y y - \beta_b i_b, \quad (5)$$

where β_b and β_y are the bond interest rate and the income elasticities of deposit demand respectively.

The real demand for goods is given by:

$$y = -\theta_l i_l - \theta_b i_b, \quad (6)$$

with θ_l and θ_b the loan interest rate and the bond interest rate elasticities of output demand respectively.

By Walras's law, we do not need to consider the bond market.

The aggregate supply function is derived from the following three equations:

$$y = a + \alpha n \quad \text{with: } \alpha \in]0; 1[, \quad (7)$$

$$p = w - a + (1 - \alpha)n, \quad (8)$$

$$w = \sigma p \quad \text{with: } \sigma \in [0; 1[, \quad (9)$$

with n labor, w wages, a total labor productivity and σ measuring the degree of nominal rigidities in the labor market. Equation (7) is a production function, (8) is a price setting equation issued from the profit maximization condition in a perfect competition framework, (9) is a wage setting equation. We assume an influence of price variations on real wages due to an imperfect adjustment of nominal wages: $\sigma < 1$. The bigger the nominal rigidities are, the smaller σ is. Using (7), (8) and (9) the aggregate supply curve can be written as:

$$y = \kappa_0 + \kappa_1 p \quad \text{with: } \kappa_0 = \frac{a}{1 - \alpha}, \kappa_1 = \frac{\alpha(1 - \sigma)}{1 - \alpha}. \quad (10)$$

Finally, given our assumption that the bond interest rate is equal to the intervention rate, i.e. $i_b = i_c$, the general equilibrium of the model is solved for four endogenous variables (y, p, i_l, r_b) using the following system of four equations:

$$\begin{cases} \text{(IS)} & y = -\theta_l i_l - \theta_b i_c, \\ \text{(LM)} & p + \beta_y y - \beta_b i_c = r_b, \\ \text{(CR)} & p - \lambda_l i_l + \lambda_b i_c + \lambda_y y = \gamma_l i_l - \gamma_b i_c + r_b, \\ \text{(AS)} & y = \kappa_0 + \kappa_1 p. \end{cases} \quad (11)$$

1.2 Comparative Statics of an Interest Rate Monetary Shock

Using (11) the comparative statics of a monetary policy shock assimilated to a change in the intervention rate can be shown to take the following form:

$$\left(\frac{dy}{di_c} \right)_a = - \frac{\theta_l(\lambda_b + \gamma_b + \beta_b) + \theta_b(\lambda_l + \gamma_l)}{\Delta}, \quad (12)$$

$$\left(\frac{dp}{di_c} \right)_a = - \frac{\theta_l(\lambda_b + \gamma_b + \beta_b) + \theta_b(\lambda_l + \gamma_l)}{\kappa_1 \Delta}, \quad (13)$$

$$\left(\frac{di_l}{di_c} \right)_a = \frac{\theta_b(\beta_y - \lambda_y) + \lambda_b + \gamma_b + \beta_b}{\Delta}, \quad (14)$$

$$\left(\frac{dr_b}{di_c} \right)_a = \left(\frac{dp}{di_c} \right)_a + \beta_y \left(\frac{dy}{di_c} \right)_a - \beta_b, \quad (15)$$

where: $\Delta = \theta_l(\lambda_y - \beta_y) + \lambda_l + \gamma_l <> 0$.

Concerning these results, there is one theoretical ambiguity linked to the value of the income elasticity of deposit demand, β_y , as compared to the income elasticity of loan demand, λ_y . If $\lambda_y > \beta_y$ then $\Delta > 0$ and there are no ambiguities relating to the sign of the income (12), price (13) and reserves

(15) multipliers, but there is instead an ambiguity concerning the sign of the interest rate multiplier (14). If, instead, $\lambda_y < \beta_y$ then $\Delta <> 0$ and the sign of all multipliers is undetermined.

Theoretically, we can solve these ambiguities directly by assuming that the interest rate multiplier is positive:

$$\left(\frac{di_l}{di_c}\right)_a > 0. \quad (H1)$$

In this case, a rise in the intervention rate will lead to a decrease in output, in prices and in banking reserves.

$$\text{If } \left(\frac{di_l}{di_c}\right)_a > 0 \Rightarrow \left(\frac{dy}{di_c}\right)_a < 0, \left(\frac{dp}{di_c}\right)_a < 0, \left(\frac{dr_b}{di_c}\right)_a < 0.$$

Empirically, provided that the model is a good description of the economy, all these ambiguities will not occur if a positive correlation is found between the intervention rate and the loan rate. In the next section, we present the empirical results indicating that the loan rate was an increasing function of the intervention rate in the period under consideration.

1.3 The Variations in the Interest Rate Spread as an Indicator of Amplification and Attenuation Effects

In order to measure the impact of the bank lending channel we need to define a standard AD/AS model as a benchmark model. This is readily done by assuming perfect substitution between bank credit and bonds. The above augmented model (11) then collapses to a model of the form:

$$\begin{cases} \text{(IS)} & y = -(\theta_l + \theta_b)i_c, \\ \text{(LM)} & p + \beta_y y - \beta_b i_c = r_b, \\ \text{(AS)} & y = \kappa_0 + \kappa_1 p. \end{cases}$$

The comparative statics results of a monetary shock in this reduced version of the model can be written as follows:

$$\left(\frac{dy}{di_c}\right)_m = -\theta_l - \theta_b < 0, \quad (16)$$

$$\left(\frac{dp}{di_c}\right)_m = -\frac{\theta_l + \theta_b}{\kappa_1} < 0, \quad (17)$$

$$\left(\frac{dr_b}{di_c}\right)_m = \left(\frac{dp}{di_c}\right)_m + \beta_y \left(\frac{dy}{di_c}\right)_m - \beta_b < 0. \quad (18)$$

From expressions (12), (16), (13) and (17) we derive the conditions under which the bank lending channel amplifies or attenuates the impact of monetary policy shocks on output and prices as compared to the traditional interest rate channel.

In the amplification case, the impact of monetary policy on output and prices is higher in the augmented model compared to that in the standard AD/AS model. Solving these inequalities indicates that this situation corresponds to an increase (decrease) in the interest rate spread between the bank

lending rate and the intervention rate in the case of a restrictive (expansionary) monetary policy.

$$\text{Amplification} \quad \left\{ \begin{array}{l} \left(\frac{dy}{di_c} \right)_a - \left(\frac{dy}{di_c} \right)_m < 0 \\ \left(\frac{dp}{di_c} \right)_a - \left(\frac{dp}{di_c} \right)_m < 0 \end{array} \right. \Rightarrow \left(\frac{di_l}{di_c} \right)_a > 1$$

In the attenuation case, the impact of monetary policy on output and prices is smaller in the augmented model compared to that in the standard AD/AS model. Solving these inequalities indicates that this situation corresponds to a decrease (increase) in the interest rate spread between the bank lending rate and the intervention rate in the case of a restrictive (expansionary) monetary policy.

$$\text{Attenuation} \quad \left\{ \begin{array}{l} \left(\frac{dy}{di_c} \right)_a - \left(\frac{dy}{di_c} \right)_m > 0 \\ \left(\frac{dp}{di_c} \right)_a - \left(\frac{dp}{di_c} \right)_m > 0 \end{array} \right. \Rightarrow \left(\frac{di_l}{di_c} \right)_a < 1$$

Finally, if the variations in income and prices are exactly the same in both frameworks, then we should observe an unchanged interest rate spread after a monetary shock.

$$\text{Neutrality} \quad \left\{ \begin{array}{l} \left(\frac{dy}{di_c} \right)_a - \left(\frac{dy}{di_c} \right)_m = 0 \\ \left(\frac{dp}{di_c} \right)_a - \left(\frac{dp}{di_c} \right)_m = 0 \end{array} \right. \Rightarrow \left(\frac{di_l}{di_c} \right)_a = 1$$

Several additional comments can be made.

First, it follows from these results and from expressions (15) and (18) that the variations in the interest rate spread are also a good indicator when distinguishing between amplification and attenuation effects of monetary policy shocks on banking reserves.

Second, a closer examination of (14) indicates that it is an increasing function of γ_b and λ_b , and a decreasing function of λ_l and γ_l . Therefore, if ceteris paribus $\gamma_l > \gamma_b$, i.e. banks are more reactive in their credit decisions to loan interest rates as compared to monetary policy-led bond interest rates, then the response of loan rates to a change in the intervention rate will be smaller and the probability of attenuation effects will increase. The same outcome will arise if ceteris paribus $\lambda_l > \lambda_b$, i.e. firms are “bank-dependent” borrowers having a more difficult access to the bond market (i.e. to credit substitutes) as compared to the loan market.

One should note that if the two main assumptions detailed in Bernanke and Blinder (1988a) apply, there will be a systematic amplification of monetary policy shocks.

$$\text{If } \lambda_y = \beta_y \text{ and if } \left\{ \begin{array}{l} \gamma_l = \gamma_b \\ \lambda_l = \lambda_b \end{array} \right. \Rightarrow \left(\frac{di_l}{di_c} \right)_a > 1.$$

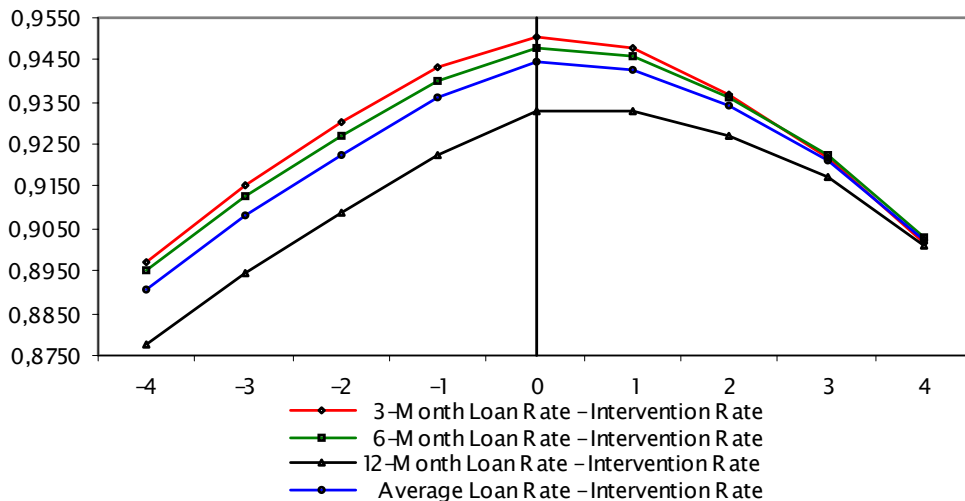
Third, Poland’s capital markets are fairly shallow: in March 2000, commercial bonds represented only 1.2 per cent and commercial papers only 5.8 per cent of total bank credit to enterprises (Łyziak, 2001). Hence, banks are almost the unique source of borrowed funds for the corporate sector. Moreover, according to the National Bank of Poland’s monthly surveys of companies, the share of firms using bank credit grew from approximately 80 per cent in 1995 to more than 85 per cent in 1999 (Łyziak, 2001). These stylized facts render attenuation effects more likely since they suggest that the value of λ_l should be strong while that of λ_b close to zero.

2 An Empirical Assessment of the Bank Lending Channel

We analyze monthly data of a sample that runs from February 1994 up to June 2001 inclusive. We construct three different interest rate spreads defined as a difference between 3-month, 6-month and 1-year minimum loan rates applied to Polish firms by major banks and the intervention rate of the central bank (cf. Appendix C for definitions).

Before a closer examination of the action of the bank lending channel in Poland, we try to verify two elements. First, we need to find out whether the loan rates and the intervention rate move in the same direction. This is the (H1) theoretical assumption of the model that must hold in order to ensure that the change in the spread makes possible a distinction between attenuation and amplification effects. Second, it is necessary to check the transmission lag between the intervention rate and the loan interest rates. To this end, we use advanced and lagged correlations between the intervention rate and the loan rates. In Figure 2 are shown the correlation coefficients between the intervention rate at date t and the loan interest rates at date $t + k$. The left-hand side of the origin shows correlations between the current monetary policy rate and lagged loan interest rates whereas the right-hand side depicts correlations between the current monetary policy rate and the future loan rates.

Figure 2. Advanced and Lagged Correlations Between Loan and Intervention Rates, II/94 - II/01



Source: National Bank of Poland and the authors' calculations.

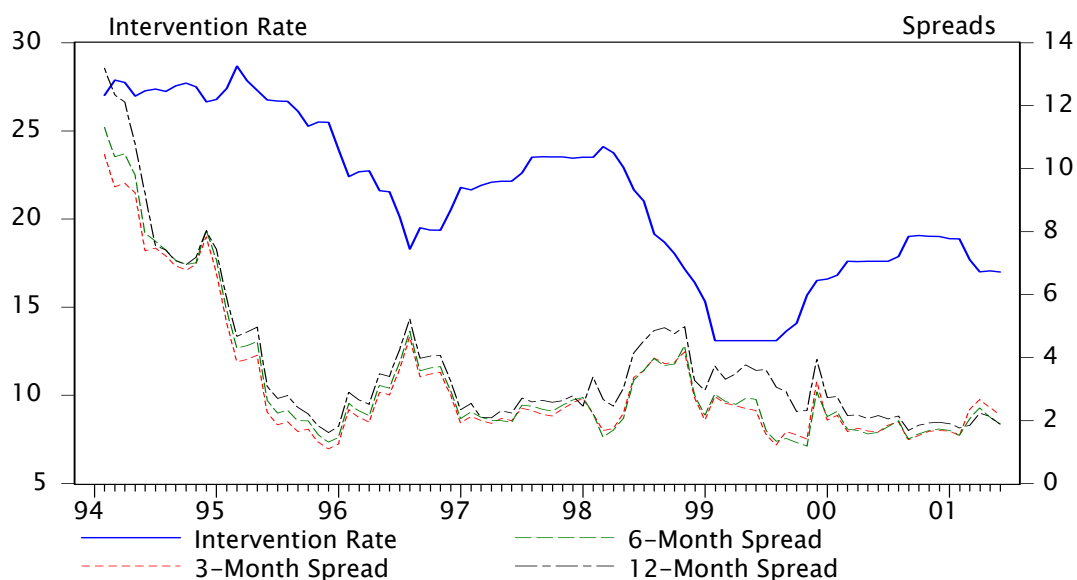
As we can see from Figure 2, there is a strong positive correlation between the intervention rate and all loan rates. Hence, the assumption (H1) of the model is satisfied.

Concerning the transmission lag, we notice that the maximum correlation is obtained for $k = 0$: this indicates an instant (within the month) pass-through of official interest rates to loan interest rates⁹.

⁹ For the 12-month loan rate the correlations for $k = 0$ and $k = 1$ are almost identical, equal to 0.93266 and 0.93261 respectively.

This finding allows us to study the reaction of the spreads calculated between current loan and intervention rates to changes in the intervention rate (see Figure 3).

Figure 3. Nominal Intervention Rate and the Interest Rate Spreads, II/94 - VI/01



Source: National Bank of Poland and the authors' calculations.

A closer look at Figure 3 shows that the period under study is not homogeneous. Recall that a scissors-like evolution of the spreads, as compared to the policy rate, indicates attenuation effects of monetary policy: the spreads decrease after a rise of the central bank's rate and rise otherwise. On the other hand, a co-movement between the spreads and the intervention rate indicates amplification effects of monetary policy: the spreads go up after an increase of the central bank's rate and diminish otherwise. Using this simple rule, we can distinguish several periods.

According to Osiński (1999), in 1994-1995, the reverse repo rate was the most important policy instrument acting on banking interest rates. Our analysis shows possible amplification effects in April and May 1994 and from June to October 1995 because of the observed co-movement between the spreads and the policy rate. Yet, care should be taken when dealing with the first period. Indeed, the important decline of all spreads resulted probably rather from the structural overliquidity of the banking system which pushed down the yield of Treasury bills (see Figure 1) and subsequently the loan interest rates, than from an enhanced monetary policy effectiveness, although there was a reduction of the intervention rate in April and May 1994. There was also an intermediate, 8-month period from October 1994 to May 1995 characterized on average by possible attenuation effects of monetary policy. However, the February 1994-December 1995 period must be analyzed with cautiousness as in several cases the average lending rate moved significantly in the opposite direction as compared to the intervention rate.

Beginning with January 1996 is a rather lengthy period, that lasted more than two years and a half, until August 1998, clearly indicating an almost systematically reduced potency of monetary policy. The important intervention rate increase since December 1996 was designed to curb the credit expansion. In September 1997, the central bank started to accept deposits directly from the public as its tightening measures did not yield the expected results on banks' behavior. Our analysis confirms the existence of a period of monetary policy weakness between December 1996 and April 1997.

The third and rather ambiguous period started in September 1998. First, we can depict several amplification episodes which occurred between September 1998 and January 1999 (except November 1998) and sporadically for example in December 1999, February 2000 and August 2000. However, banks were clearly reducing the impact of monetary shocks for example in February 1999, March and September 2000 and in the March-April 2001 period.

3 A Simple Regime-Switching Model

Let us consider the following model:

$$\dot{D}_t = X'_{1,t}\beta_1 + \varepsilon_{1,t}, \quad (19)$$

$$\dot{S}_t = X'_{2,t}\beta_2 + \varepsilon_{2,t}, \quad (20)$$

$$\dot{Q}_t = \text{Min} \left(\dot{D}_t, \dot{S}_t \right), \quad (21)$$

where \dot{D}_t denotes the annual growth rate of the demanded quantity of loans during period t , \dot{S}_t the annual growth rate of the supplied quantity of loans during period t . $X'_{1,t}$ and $X'_{2,t}$ denote the variables that influence \dot{D}_t and \dot{S}_t respectively, and $\varepsilon_{1,t}$ and $\varepsilon_{2,t}$ are the residuals. \dot{Q}_t is the actual annual growth rate of the observed quantity in the loan market during period t . The above model can be considered as a regime-switching model that allows for two regimes for characterizing \dot{Q}_t . Given equation (21), the growth rate of the amount of loans exchanged in the market corresponds to the minimum of the loan supply and demand growth rates. In other words, a demand (supply) regime takes place if the growth rate of the quantity of loans is determined by the variables and their parameters associated with the annual increase in loan demand (supply). The occurrence of regimes, i.e. any divergence between \dot{D}_t and \dot{S}_t , indicates (with a 12-month lag) the existence of a disequilibrium in level in the loan market. Indeed, if the level of demand is equal to the level of supply in each date (i.e. $D_t = S_t$), this implies an equality of the annual growth rates of loan demand and supply (i.e. $\dot{D}_t = \dot{S}_t$). The disequilibrium is the result of market imperfections leading to an incomplete price adjustment of the loan interest rate and therefore to a possible distortion of the monetary policy impulses. In the case of credit rationing, the speed and effectiveness of monetary contractions is substantially increased (Tucker, 1968). At the same time, the efficiency of the monetary policy is not symmetric since following a

monetary expansion, there may be substantial lags and the overall impact may be less than desired if there is an excess demand for bank loans (Sealey, 1979). However, since the quantity of loans and therefore the estimates of the unobservable loan demand and supply variables are defined in the paper as annual growth rates, our methodology precludes the identification of the type of disequilibrium in level (whether the level of loan demand exceeds the level of loan supply or vice versa). Yet, we should observe a supply (demand) regime if the level of the loan quantity is equal to the level of the loan supply (demand).

As shown by Maddala and Nelson (1974), with condition (21), the model itself determines the probabilities with which each observation belongs to either \dot{D}_t or \dot{S}_t . In what follows, we briefly develop the theoretical underpinnings of this result. We also discuss the choice of initial conditions which is of primary importance in order to get consistent estimators of the structural parameters of the model.

We assume that both innovations in each function are *i.i.d.* Gaussian processes.

Assumption H_1 We assume that $\varepsilon_t = (\varepsilon_{1,t} \varepsilon_{2,t})'$ is a *i.i.d.* vector and normally distributed $N(0, \Omega)$, with:

$$\Omega = E(\varepsilon_t \varepsilon_t') = \begin{pmatrix} \sigma_1^2 & \sigma_{12} \\ \sigma_{12} & \sigma_2^2 \end{pmatrix}. \quad (22)$$

3.1 *ML* Estimation of Parameters

Let θ denote the vector of structural parameters $\theta = (\beta_1 \beta_2 \sigma_1 \sigma_2 \sigma_{12})'$. In order to propose an *ML* estimate of θ , we have to compute the marginal density, denoted $f_{\dot{Q}_t}(\dot{q}_t)$, of the only observable variable \dot{Q}_t . For that, we first consider the joint density of \dot{D}_t and \dot{S}_t , denoted $g_{\dot{D}_t, \dot{S}_t}(\dot{d}_t, \dot{s}_t)$. Given the definition of the disequilibrium in the model, we know that:

$$f_{\dot{Q}_t}(\dot{q}_t) = f_{\dot{Q}_t | \dot{D}_t < \dot{S}_t}(\dot{q}_t) + f_{\dot{Q}_t | \dot{S}_t < \dot{D}_t}(\dot{q}_t). \quad (23)$$

Then, we get the corresponding marginal density of \dot{Q}_t on the two subsets (cf. Appendix A):

$$f_{\dot{Q}_t | \dot{D}_t < \dot{S}_t}(\dot{q}_t) = \int_{\dot{q}_t = \dot{d}_t}^{\infty} g_{\dot{D}_t, \dot{S}_t}(\dot{d}_t, z) dz, \quad (24)$$

$$f_{\dot{Q}_t | \dot{S}_t < \dot{D}_t}(\dot{q}_t) = \int_{\dot{q}_t = \dot{s}_t}^{\infty} g_{\dot{D}_t, \dot{S}_t}(z, \dot{s}_t) dz. \quad (25)$$

Finally, we get the unconditional density function of \dot{Q}_t as:

$$f_{\dot{Q}_t}(\dot{q}_t) = f_{\dot{Q}_t}(\dot{q}_t, \theta) = \int_{\dot{q}_t}^{\infty} g_{\dot{D}_t, \dot{S}_t}(\dot{q}_t, z) dz + \int_{\dot{q}_t}^{\infty} g_{\dot{D}_t, \dot{S}_t}(z, \dot{q}_t) dz. \quad (26)$$

Next, conditionally to a structural parameters set θ and a sample of observable variables \dot{Q}_t , $X_{1,t}$ and $X_{2,t}$ observed on T periods, the log-likelihood function of the model is then defined by:

$$L(\theta) = \sum_{t=1}^T \log [f_{\dot{Q}_t}(\dot{q}_t, \theta)]. \quad (27)$$

If we assume that both residuals ε_1 and ε_2 are independent ($\sigma_{12} = 0$), the unconditional density function of \dot{Q}_t can be expressed as follows:

$$f_{\dot{Q}_t}(\dot{q}_t) = \frac{1}{\sigma_1} \phi\left(\frac{x'_{1,t}\beta_1 - \dot{q}_t}{\sigma_1}\right) \Phi\left(\frac{x'_{2,t}\beta_2 - \dot{q}_t}{\sigma_2}\right) + \frac{1}{\sigma_2} \phi\left(\frac{x'_{2,t}\beta_2 - \dot{q}_t}{\sigma_2}\right) \Phi\left(\frac{x'_{1,t}\beta_1 - \dot{q}_t}{\sigma_1}\right), \quad (28)$$

where $\phi(\cdot)$ denotes the normal $N(0, 1)$ density function and $\Phi(\cdot)$ the corresponding cumulative distribution function. The proof is done in Appendix B. In this case, the first and second order derivatives of $L(\theta)$ can be computed analytically (Maddala and Nelson, 1974). Then, we can use an iterative procedure like the Newton-Raphson procedure to obtain the ML estimates of the structural parameters. However, this method should be used very carefully, since the likelihood may be not convergent, whether we choose one or both variance residual σ_i , close to 0 (Gouriéroux, 1989). For instance, let us consider positive finite values for β_1, β_2 and σ_2 . In this case, the first member of (28), i.e.

$$\frac{1}{\sigma_1} \phi\left(\frac{x'_{1,t}\beta_1 - \dot{q}_t}{\sigma_1}\right) \Phi\left(\frac{x'_{2,t}\beta_2 - \dot{q}_t}{\sigma_2}\right),$$

is finite. But the first term is degenerated if σ_1 tends to zero. Indeed, if σ_1 tends to zero, $1/\sigma_1\phi(\cdot)$ tends to ∞ , since given the properties of the normal distribution, $\lim_{x \rightarrow \infty} x\phi(x) = \infty$. The term $\Phi(\cdot)$ tends to 1, since $\lim_{x \rightarrow \infty} \Phi(x) = 1$. By analogy, the same analysis applies to the second member of (28) which is degenerated if σ_2 tends to zero for given positive finite values of β_1, β_2 and σ_1 . To summarize, the global maximum of the likelihood function is infinity, if one of the residual variances (or both) tends to zero. Therefore, the application of the standard ML procedure must be adapted in this case, as only local maximum must be searched for.

3.2 Initial Conditions

There are various methods to obtain the initial conditions on structural parameters θ in the ML iteration. Here, we use a two step OLS procedure. First, we consider the linear regressions of the observation \dot{q}_t on the exogenous variables sets in both functions: $\dot{q}_t = x'_{i,t}\hat{\gamma}_i + \mu_{i,t}$, with $i = 1, 2$. Given the realizations of $\hat{\gamma}_1$ and $\hat{\gamma}_2$, we compute a first approximation of demand and supply growth rates, as $\tilde{d}_t = x'_{1,t}\hat{\gamma}_1$ and $\tilde{s}_t = x'_{2,t}\hat{\gamma}_2$. Even if we know that $\hat{\gamma}_1$ and $\hat{\gamma}_2$ are not convergent estimators of β_1 and β_2 , we built two subgroups of observations. In the first subgroup, denoted by index d , we consider only the observations on $\dot{Q}_t, X_{1,t}$ and $X_{2,t}$ for which we have $\tilde{d}_t \leq \tilde{s}_t$. In the second subgroup, we consider the observations for which we have $\tilde{s}_t \leq \tilde{d}_t$. The second step of the procedure consists in applying the OLS on both subgroups:

$$\dot{q}_t^{(d)} = x'_{1,t} \tilde{\beta}_1 + \tilde{\mu}_{1,t} \quad \text{and} \quad \dot{q}_t^{(s)} = x'_{2,t} \tilde{\beta}_2 + \tilde{\mu}_{2,t}. \quad (29)$$

Then, we use the *OLS* estimates $\tilde{\beta}_i$ as starting values for β_i in the *ML* iteration. For the parameters σ_1 and σ_2 , we adopt the following starting values:

$$\tilde{\sigma}_i = \frac{1}{n_i} \sum_{j=1}^{n_i} \tilde{\mu}_{i,j} \quad i = 1, 2, \quad (30)$$

where n_1 denotes the size of the “demand growth” subgroup of observations for which we have $\tilde{d}_t \leq \tilde{s}_t$, and n_2 denotes the size of the corresponding “supply growth” subgroup.

In the following experiment, we can observe that such a procedure gives starting values very close to the true solution θ_0 . We consider 10 000 replications of the model (21) with the following parameters: $\beta_1 = (\beta_{1,1}, \beta_{1,2})' = (5, 2)'$, $\beta_2 = 3$, $\sigma_1 = \sigma_2 = 0.5$, $T = 100$. For each simulation, the realizations of the scalar components of the exogenous variables X_1 and X_2 are drawn in normal distributions $N(0, \sigma_x^2)$, with $\sigma_x = 0.3$. The results of this experiment are reported in Table 2.

Table 2. Results of the Monte Carlo Experiment on Starting Values

Parameter	True Value	One Step OLS	Two Step OLS
$\beta_{1,1}$	5	2.5007 (0.5192)	4.9882 (0.2871)
$\beta_{1,2}$	2	0.9990 (0.4059)	1.9954 (0.2971)
β_2	3	1.5023 (0.4958)	2.9984 (0.3120)
σ_1	0.5	0.8207 (0.0663)	0.4754 (0.0500)
σ_2	0.5	1.0576 (0.1006)	0.4807 (0.0511)

Figures in parentheses are standard errors.

We can verify that one step *OLS* are not convergent in this regime-switching model. In particular, given our simulation parameters, it can be shown that both regimes have the same theoretical unconditional probability to appear. This implies that the one step *OLS* estimators of β_1 and β_2 converges to $\beta_1/2$ and $\beta_2/2$. These estimates could be considered as starting values of the *ML* iteration if the optimization algorithm is powerful enough to converge to the true solution θ_0 . However, in our simulations, the two step *OLS* always gives a better set of starting values. Indeed, the mean of the realizations of $\tilde{\beta}_i$ and $\tilde{\sigma}_i$ are very close to the true values of parameters, and the difference is inferior to 10^{-2} for the β_i . Hence, in the application we will use the realizations of the two step *OLS* estimators as starting values for *ML* iterations.

3.3 Probability of Both Regimes

Given the estimated values of the parameters, we can compute the probability that the observation \dot{q}_t belongs to the demand or the supply regime. Let us assume that structural parameters

$\theta = (\beta_1 \beta_2 \sigma_1 \sigma_2)'$ are known. The probability to be in the demand regime, at time t , is given by:

$$\begin{aligned} P(\dot{D}_t < \dot{S}_t) &= P(X'_{1,t}\beta_1 + \varepsilon_{1,t} < X'_{2,t}\beta_2 + \varepsilon_{2,t}) \\ &= P\left(\frac{\varepsilon_{1,t} - \varepsilon_{2,t}}{\sigma} < \frac{X'_{2,t}\beta_2 - X'_{1,t}\beta_1}{\sigma}\right). \end{aligned}$$

Under general assumption H_1 , the transformed variable $\varepsilon_{1,t} - \varepsilon_{2,t}$ is normally distributed with a variance equal to $\sigma^2 = \sigma_1^2 + \sigma_2^2$. Then, the reduced variable $(\varepsilon_{1,t} - \varepsilon_{2,t})/\sigma$ follows a $N(0, 1)$. Subsequently, the probability that the observation \dot{q}_t belongs to the demand regime, denoted $\pi_t^{(d)}$, can be computed as the corresponding $N(0, 1)$ fractile:

$$\pi_t^{(d)} = P(\dot{D}_t < \dot{S}_t) = \Phi(h_t) = \int_{-\infty}^{h_t} \frac{1}{\sqrt{2\pi}} e^{-\frac{x^2}{2}} dx, \quad (31)$$

where $h_t = (X'_{2,t}\beta_2 - X'_{1,t}\beta_1)/\sigma$, and $\Phi(h_t)$ denotes the normal cumulative distribution function. Symmetrically, the probability to be in the supply regime, denoted $\pi_t^{(s)}$, is defined by $P(\dot{S}_t < \dot{D}_t) = 1 - \Phi(h_t)$.

If we have convergent estimators of structural parameters θ , we can replace these estimators into expression (31) to obtain an estimate of the threshold h_t and an estimate of the probability of the demand regime. Let us consider three convergent estimators $\hat{\beta}_1, \hat{\beta}_2$ and $\hat{\sigma}$. We define the estimate \hat{h}_t as $\hat{h}_t = (X'_{2,t}\hat{\beta}_2 - X'_{1,t}\hat{\beta}_1)/\hat{\sigma}$, and the estimate of $\pi_t^{(d)}$ is defined by¹⁰:

$$\hat{\pi}_t^{(d)} = \Phi(\hat{h}_t). \quad (32)$$

4 An Empirical Assessment of the Loan Market Disequilibrium

Before estimating the regime-switching model, we review a few stylized facts of the Polish loan market as a first approach.

First, in the early nineties, the stability of the banking sector was affected by the so-called “bad loans” problem whose origin can basically be found in the 1990-1991 transformational recession. The introduction of new definitions of classified loans in November 1992 revealed a 30 per cent ratio of non-performing loans to total loans. For this reason, the authorities adopted a comprehensive framework aiming at restoring the foundations of a strong and healthy financial intermediation. *The Law on Financial Restructuring of State Enterprises and Banks* became effective in March 1993. It introduced a decentralized approach to solve the problem: banks were obliged to restructure their bad loans portfolio so as to be involved in the recapitalization programme¹¹. Although 1993 is often

¹⁰ Besides, we can compute an interval of confidence for $\hat{\pi}_t^{(d)}$ as $IC_\alpha = [\Phi(\hat{h}_t^{\text{inf}}), \Phi(\hat{h}_t^{\text{sup}})]$, where \hat{h}_t^{inf} and \hat{h}_t^{sup} are the extremum of the interval, such that, for a risk of $\alpha\%$ and for non stochastic regressor: $P[\Phi(\hat{h}_t^{\text{inf}}) < \hat{\pi}_t^{(d)} < \Phi(\hat{h}_t^{\text{sup}})] = 1 - \alpha$.

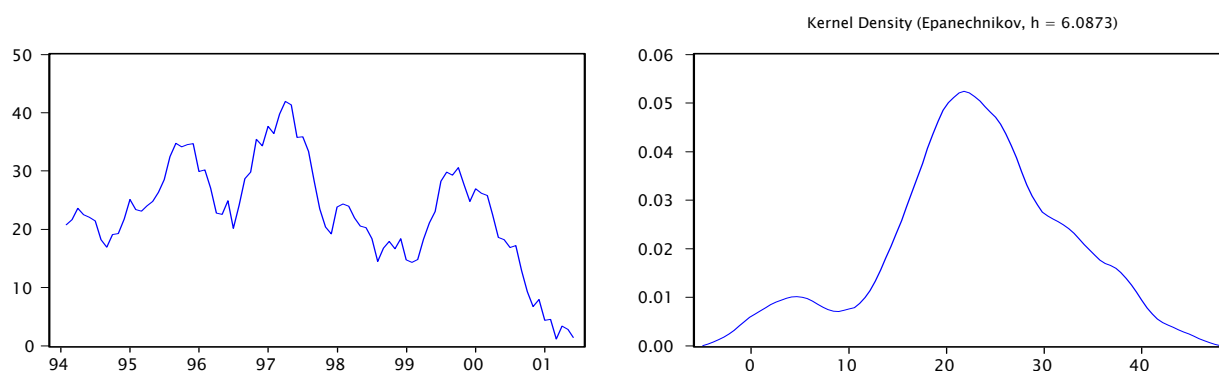
¹¹ Taking the average current exchange rate for each year, the governmental recapitalization amounted to 1.14 USD billion in 1993 and 1.75 USD billion in 1994.

considered as a typical “credit crunch” year, the interesting question to be addressed is whether this state of the market lasted in the subsequent years. What we know for sure is that the end-of-year share of non-performing loans to total loans for enterprises and households decreased steadily from 31.2 per cent in 1993 to 28.3 per cent in 1994 reaching a low point of 9.8 per cent in September 1997.

Second, fixed exchange rate policies followed until at least the end of July 1998¹² can be expected to have had a long-standing impact on the loan market. Sterilization operations of capital inflows created a structural overliquidity of the banking system, defined as a net indebtedness of the central bank towards commercial banks. This has been a permanent situation since the end of 1993: it potentially renders demand regimes more likely. At the same time, we can also expect that imperfectly sterilized capital inflows had a direct positive impact on the annual rate of growth of loan supply.

Third, the instability of real activity increased since the last quarter of 1998. Whereas the average annual growth rate of GDP was about 5.6 per cent in the 1993-1998 period, it amounted to 4.1 per cent in 1999 and started to fall below trend from 5.9 per cent year-on-year in the first quarter of 2000 to 2.4 per cent in the fourth quarter. In the second quarter of 2001 it was 0.9 per cent. Simultaneously, the share of non-performing loans to total loans experienced an upsurge from 10.7 per cent in December 1998 to 13.2 per cent in December 1999 and attained a 16.5 per cent level in June 2001 (for the corporate sector only, the corresponding figures were 11.9, 15.1 and 18.6 per cent respectively). These stylized facts had probably a negative impact on the growth rates of loan demand and supply.

Figure 4. Nominal Growth Rate of Loans to the Corporate Sector and its Empirical Density, II/94 - VI/01



Source: National Bank of Poland and the authors' calculations.

Figure 4 represents the nominal growth rate of the loan series used in the study and its corresponding Kernel density. The empirical density has the general form of a mixture of normal distributions. This observation, with the history of the Polish loan market outlined above, renders relevant the disequilibrium assumption and, therefore, the estimation of the regime-switching model

¹² According to Polański (2000), date at which the central bank stopped its direct interventions on the foreign exchange market.

presented in the previous section.

4.1 Data

The data set covers the period from February 1994 to June 2001, including 89 monthly observations. All data were obtained from the National Bank of Poland, except the industrial production which comes from the Central Statistical Office (GUS). As to the loan series, we used total zloty denominated loans up to one year extended to resident and non resident firms. This series represents, on average for the period under consideration, 36.7 per cent of total credit extended to firms for the sample of banks used in the study (cf. Appendix C for definition) and 27.6 per cent taking the whole Polish banking sector. In all models, the variables are expressed as percentage annual growth rates, except for the interest rates which are measured in percentage points. Unless otherwise indicated, all variables are in nominal terms.

4.2 Specification Research

We consider several models. The first one (model 1) is based on the theoretical model of the bank lending channel presented in section 1. It is defined by:

$$\begin{aligned}\dot{D}_t &= \beta_0 + \beta_1 IL_t + \beta_2 IC_t + \beta_3 PROD_t + \varepsilon_{1,t}, \\ \dot{S}_t &= \alpha_0 + \alpha_1 IL_t + \alpha_2 IC_t + \alpha_3 DEP_t + \varepsilon_{2,t}, \\ \dot{Q}_t &= \text{Min}(\dot{D}_t, \dot{S}_t),\end{aligned}\tag{Model 1}$$

with $\varepsilon_{i,t}$ *i.i.d.* $N(0, \sigma_i^2)$ and where \dot{D}_t denotes the annual growth rate of the demand for bank loans, \dot{S}_t the annual growth rate of the supply of bank loans and \dot{Q}_t the annual growth rate of the observed amount of loans. The exogenous variables of both loan demand and supply functions and their expected signs are as follows.

The loan interest rate, IL_t , is defined as an arithmetical mean of 3-month, 6-month and 1-year weighted averages for minimum loan rates applied to Polish firms by major banks. In accordance with the model, it is expected to be statistically significant with a negative sign in the demand function and a positive sign in the supply one. IC_t is the intervention rate of the NBP (cf. Appendix C for definition). Due to the substitution effect, IC_t should have the opposite sign to IL_t in the two equations. In the disequilibrium loan market literature¹³, a lagged index of industrial production is often used to approximate the firms' and the banks' expectations about future economic activity and to have a positive sign. Following the Bernanke and Blinder (1988a,b) and the bank lending channel literature, we assumed in the model a positive dependence of loan demand on output invoking working

¹³ See, for instance, Sealey (1979), Kim (1999), Pazarbaşıoğlu (1997).

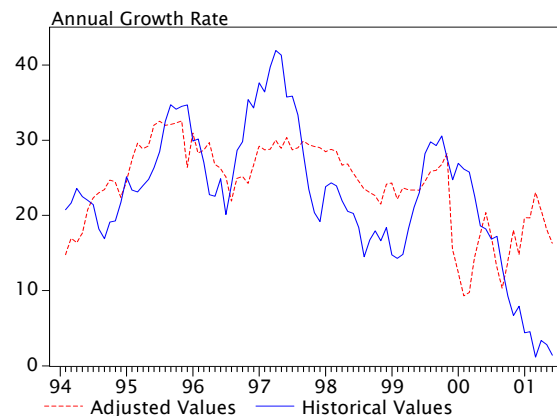
capital or liquidity considerations. Yet, the latter assumption is rather ambiguous if the corporate sector loan demand is considered. As a matter of fact, a drop in industrial production will probably strengthen the liquidity constraint of firms, thus increasing their short-term credit demand. Putting all these arguments together, the theoretical sign of the parameter of industrial production, $PROD_t$, is indeterminate¹⁴. DEP_t denotes the total zloty and foreign currency deposits of firms and households of our sample of banks and is expected to have a positive coefficient.

The results are reported in Table 3. Except the loan interest rate in the loan supply equation, all coefficients in both functions have the expected signs and the industrial production is positively signed. However, as measured by the asymptotic t-statistics values, only interest rates in the loan demand function are statistically significant. Moreover, the quality of the model is fairly low as the R-squared and the adjusted R-squared statistics indicate. The visual inspection simply depicts the latter outcome: the annual growth rate of loans issued from the estimated model match rather poorly the historical values. But encouragingly, the estimated demand and supply regimes' periods (not reported here) correspond approximately with the ones (reported later on) that we found using a more sophisticated version of the model.

Table 3. ML Estimates and the Adjustment on Quantities of Model 1

Variable	Demand Function	Supply Function
Constant	23.61 (3.34)	24.33 (0.55)
IL_t	-1.87 (-4.20)	-1.66 (-0.27)
IC_t	2.22 (3.31)	-0.61 (-0.12)
$PROD_t$	0.09 (0.67)	—
DEP_t	—	2.31 (1.35)
σ_i^2	5.88	11.49
$R^2 = 0.26$	$\bar{R}^2 = 0.20$	$L = -295.80$

Asymptotic t-statistics are in parentheses, L: log-likelihood.



The drawback of estimating model 1 is that there is only one variable specific to each function, namely the industrial production on the demand side and deposits on the supply side. Therefore, in order to better discriminate between the growth rates of loan demand and supply, we estimated a second model (model 2) including three additional variables as compared to model 1. The first one, denoted SAV_t , is a proxy for firms' cash flow. It is equal to the total time deposits of the firms and is expected to have a negative impact on loan demand. The purpose of the second variable, called OL_t , is to capture the impact of other loans extended to the corporate sector. It is defined as the following ratio:

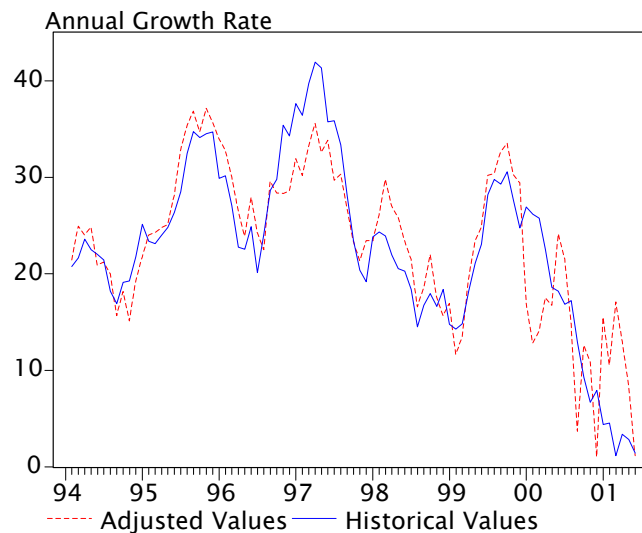
¹⁴ It should be noted that a possible negative sign of the output coefficient in the loan demand function does not affect the interest rate spread indicator, if the assumption (H1) of the model holds.

sum of zloty denominated loans beyond one year and of total foreign currency loans divided by total extended loans. Introduced in the loan demand equation, the corresponding coefficient is expected to have a positive sign in the case of complementary effects or a negative one if other banking means can be substituted for zloty denominated loans up to one year. Finally, the third variable, denoted $LGOR_t$, measures the net liquidity impact on banks of the foreign reserves accumulation by the NBP. $LGOR_t = GOR_t - ACB_t - ATB_t$, with GOR_t the gross official reserves of the NBP, ACB_t the overall value of banking funds absorbed by sterilization operations led by the NBP (including the net value of reverse repo, outright and NBP bills issue operations and the value of required reserves) whereas ATB_t represents the value of banking assets invested in Treasury bills. One could claim that $LGOR_t$ is no longer an appropriate indicator since the implementation of the floating exchange rate system, that is, since at least June 1999 (date at which the NBP definitely stopped its foreign exchange market interventions by abolishing fixing transactions). However, according to OECD (2001), in 2000 and early 2001, the NBP continued to make off market foreign exchange transactions as an agent of the government, which ultimately had the same impact on banking liquidity as market interventions.

Table 4. ML Estimates and the Adjustment on Quantities of Model 2

Variable	Demand Function	Supply Function
Constant	31.62 (5.71)	6.05 (0.17)
IL_t	-1.02 (-3.06)	2.50 (0.51)
IC_t	1.44 (3.13)	-4.51 (-1.13)
$PROD_t$	0.05 (0.56)	—
DEP_t	—	2.34 (3.35)
$LGOR_t$	—	0.37 (3.00)
SAV_t	-0.26 (-5.36)	—
OL_t	-1.77 (-5.17)	—
σ_i^2	3.34	7.28
$R^2 = 0.72$		$\bar{R}^2 = 0.68$
$L = -250.43$		

Asymptotic t-statistics are in parentheses, L: log-likelihood.



The results of model 2 estimates are reported in Table 4. All variables have the expected sign and, except the interest rates in the loan supply equation and the industrial production, are significant. In addition, the quality of the model increased markedly as is proved by the R-squared statistics, the log-likelihood and the adjusted values of loans.

In model 2, we assumed that all exogenous variables affect contemporaneously the loan demand and supply functions. In fact, we do not know economically what kind of structure of lags is underlying the transmission mechanism between our set of exogenous variables and the annual growth rate of

actual loans. Thus, going one step further, we establish the optimal lag structure and thus the best quality of the model using purely statistical criterions. More precisely, we test a specification 1 for which we impose lags on all variables versus a specification 2 with lagged interest rates only. The results are reported in Table 5.

Table 5. Optimal Lag Tests

Number of lags	lags on all variables specification 1		lags on interest rates only specification 2	
	AIC	SC	AIC	SC
3	317.02	348.93	272.43	304.34
2	311.39	343.44	277.08	309.14
1	302.53	334.73	270.90	303.11
0	302.86	335.22	302.86	335.22

AIC: Akaike information criterion; SC: Schwarz criterion.

Whatever the number of lags, specification 2 dominates specification 1 as it yields systematically lower values for both criterions and is thus a preferred outcome. For specification 2, the minimum of both criterions is reached by using one lag structure on interest rates.

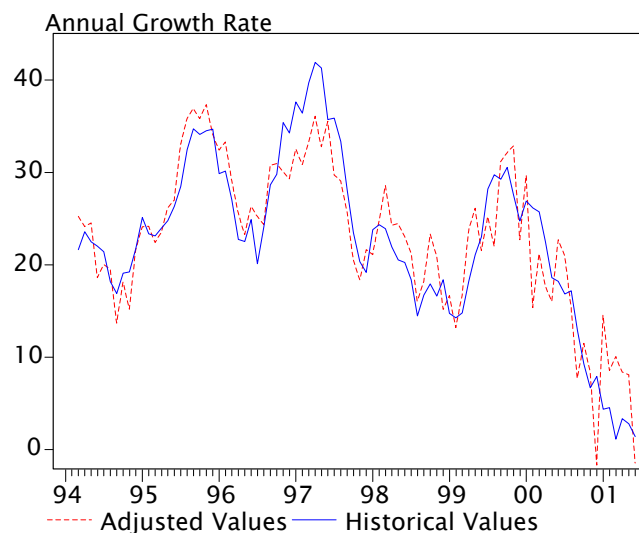
4.3 The Final Specification

The results of the final specification (model 3) are shown in Table 6.

Table 6. ML Estimates and the Adjustment on Quantities of Model 3

Variable	Demand Function	Supply Function
Constant	46.27 (9.05)	-37.91 (-2.68)
IL_{t-1}	-0.78 (-3.00)	9.77 (3.47)
IC_{t-1}	0.44 (1.15)	-9.75 (-3.79)
$PROD_t$	0.20 (2.26)	—
DEP_t	—	1.85 (7.45)
$LGOR_t$	—	0.46 (6.60)
SAV_t	-0.28 (-6.15)	—
OL_t	-2.27 (-6.45)	—
σ_i^2	3.33	5.44
$R^2 = 0.80 \quad \bar{R}^2 = 0.77 \quad L = -243.45$		

Asymptotic t-statistics are in parentheses, L : log-likelihood.



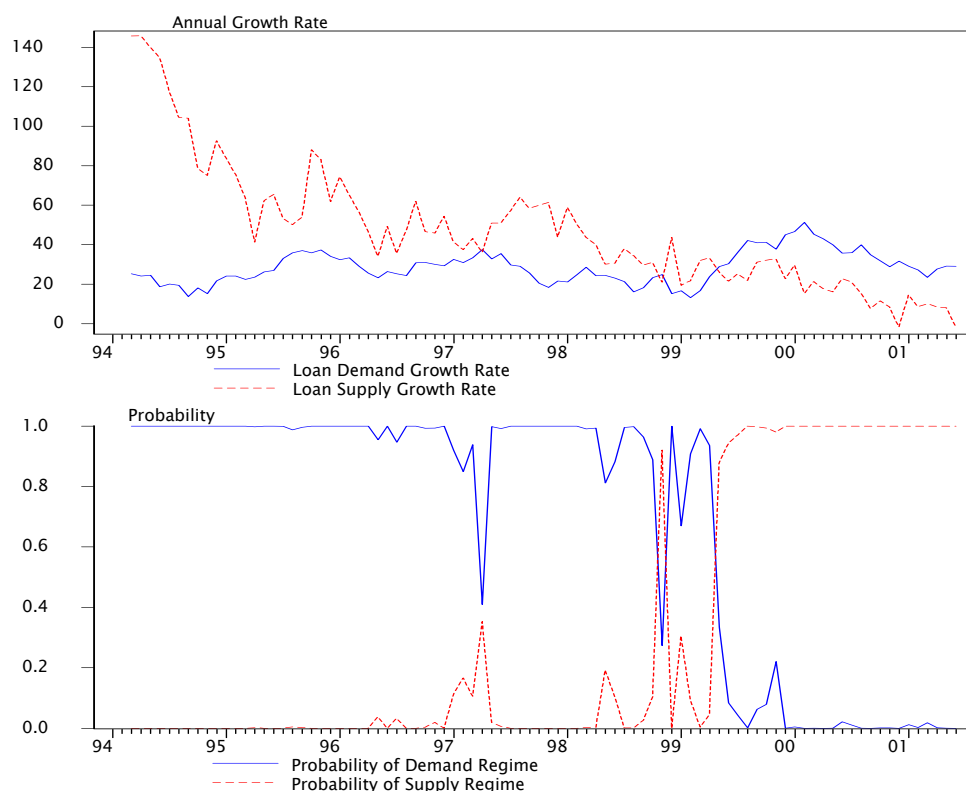
As expected, the values of R^2 , \bar{R}^2 and L are higher than in model 2 and the estimated quantities match more closely the actual ones, particularly at the end of the period. As in the previous models, the industrial production has a positive sign in the loan demand function but the value of the associated parameter is very low. Moreover, OL_t has a statistically significant coefficient and is negatively signed,

indicating that other banking loans can be substituted for the zloty denominated loans up to one year. All other variables have the expected signs and, except for the intervention rate in the loan demand equation, have statistically significant parameters in the two equations. We note that the interest rates parameters in the loan supply equation are considerably higher than those in the loan demand function. This suggests that banks are much more sensitive to interest rates in their loans' decisions compared to the corporate sector. The high t-statistic of $LGOR_t$ shows that, despite huge sterilization operations and the growth of Treasury bills in banking assets, the heavy build-up of the NBP's foreign exchange reserves had a strong positive impact on the loan supply growth. However, the estimated coefficient is much lower than the one associated with DEP_t . This indicates that the total deposits of the non-bank private sector ultimately had a stronger effect on banks' loan supply behavior.

Figure 5 shows the fitted loan demand and supply growth rates and the corresponding probabilities of each regime. Recall that the growth rate of the quantity of loans exchanged in the market corresponds to the minimum of the loan supply and demand growth rates. Put differently, a demand (supply) regime occurs if the growth rate of loans is determined by the variables and their parameters associated with the annual increase in loan demand (supply). The results clearly reveal the existence of two regimes. The loan market was characterized by a strong demand regime until at least the end of 1997 and by a supply regime since May 1999. In any case, since the loan demand and supply growth rates are almost never identical, the implication is that disequilibrium is, in fact, a permanent characteristic of the loan market since 1994.

The period under consideration begins with an outstanding, almost 150 per cent annual growth rate of the loan supply. The reliability of this result will be shown later on. Since then, the growth rate of the loan supply steadily declined to approximately 43 per cent in April 1995, with a growth rate differential between loan supply and demand amounting to 20 per cent. Interestingly, in this downward trend, there was a temporary 17 per cent increase in December 1994, probably linked to the denomination of the Polish zloty, which started at the beginning of 1995. Subsequently, despite an important increase in the loan demand growth between April and November 1995, there was a simultaneous surge in the loan supply growth leading to a stronger demand regime. This situation was probably brought about by the imperfectly sterilized huge foreign assets growth: in 1995, the stock of gross official reserves increased from 6 to 15 USD billion. As of December 1996, in order to slow down the demand of goods, the NBP started to tighten up progressively its monetary policy, rising both the interest rates and the reserve requirement ratios. However, considering that these measures did not yield the expected results on banks' behavior, the central bank started to accept deposits directly from the public in mid-September 1997. Our results bring some ex-post justification for this unprecedented action, as far as the zloty up to one year loan market to the corporate sector is considered.

Figure 5. Estimated Loan Demand and Supply Growth Rates and the Probabilities of Each Regime, II/94-VI/01



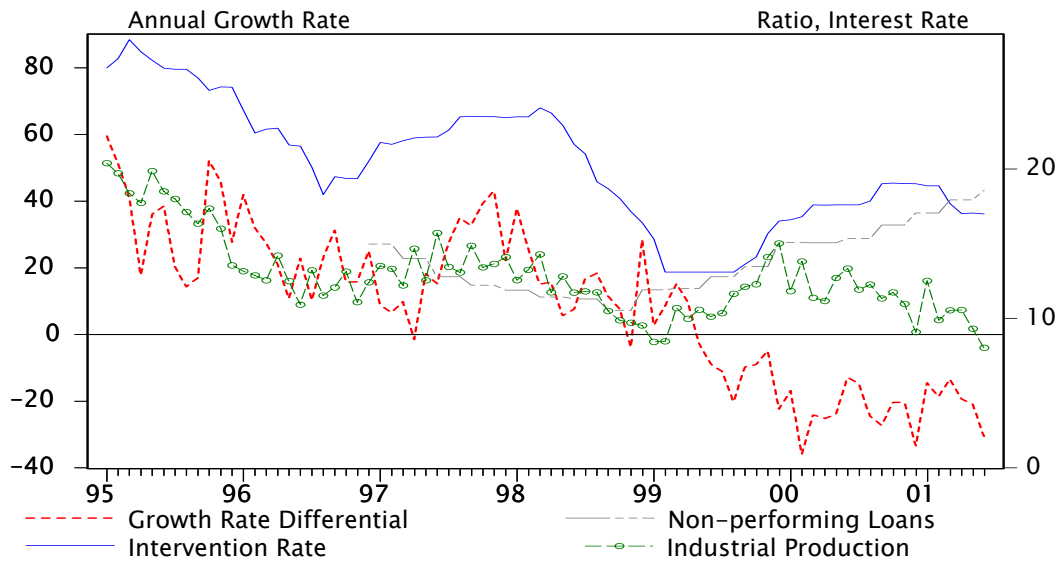
Since December 1996, the discrepancy between loan demand and supply growth rates actually decreased. However, after a turning point in April 1997, it started to rise at an increasingly fast rate reaching an all time high in July 1997. Overall, one should note that since that time the loan supply growth follows a downward trend. The growth rate differential between loan supply and demand was progressively reduced to almost zero in November 1998. The supply regime appeared in May 1999. However, the discrepancy between loan supply and loan demand growth rates increased markedly and the market definitely switched into a supply regime in December 1999. In the same month, there was also a negative reversal in the industrial production growth. Since February 2000, the loan market was characterized by a downward falling supply and demand growth rates.

Figure 6 plots the intervention rate, the annual growth rate differential in the loan market (growth rate of supply – growth rate of demand), the annual increase in industrial production and the ratio of non-performing loans to total loans for the corporate sector since 1995¹⁵. Several comments can be made.

First, it is interesting to note the existence of an important co-movement between the growth rate of industrial production and the growth rate differential between loan supply and demand. This could

¹⁵ We use the ratio for the entire banking sector only because we had neither the corresponding data for our loan series nor for our sample of banks.

Figure 6. Growth Rate Differential Between Loan Supply and Demand and Economic Conditions, I/95 - VI/01



Source: National Bank of Poland, Central Statistical Office (GUS) and the authors' calculations.

mean that banks have adjusted their loan supply growth to the observed output.

Second, the ratio of non-performing loans started to rise in December 1998, that is nine months before the beginning of a new stage of monetary tightening and exactly five months before the first signs of the supply regime. However, the quality of extended loans continued to deteriorate when the supply regime occurred and the intervention rate was going up.

Third, it is difficult to say what kind of shock brought about a regime switch in the loan market. This issue must be analyzed cautiously. Nevertheless, the almost simultaneity of a supply regime, of monetary stringency and of a decline in output is striking. At any rate, a coexistence of a raising monetary policy rate and of a supply regime means that banks were probably amplifying an increasingly restrictive monetary policy.

4.4 The Robustness of the Final Specification

We performed several tests in order to check the robustness of our results.

First, we controlled for the existing high collinearity of IL_{t-1} with IC_{t-1} by using the spread $IL_{t-1} - IC_{t-1}$ and the intervention rate IC_{t-1} instead of IL_{t-1} and IC_{t-1} in each equation. The results reported in Table 8 (cf. Appendix C), are identical to those of model 3.

Second, we estimated the model using CPI (cf. Appendix D) and PPI (cf. Appendix E) deflated data, but considering the real industrial production as is calculated by the Central Statistical Office (GUS). The results are basically unchanged compared to the nominal specification. In both cases, there was a period of a strong demand regime until the end of 1997, and a period of a marked supply

regime since December 1999.

Third, we checked for the nominal as well as for the real specifications that our findings are not sensitive to the suppression of any of the three additional variables SAV_t , OL_t and $LGOR_t$.

Fourth, taking three different starting dates (January 1995, 1996 and 1997), we did not find a major temporal stability problem in our results. The model generated similar to baseline estimates of loan demand and supply growth rates in nominal as well as in real terms, although sometimes one or two interest rates or the OL_t coefficients had the “wrong” sign or/and were not significant.

Fifth, we can justify the reliability of the impressive supply growth rate in 1994 using two arguments, a statistical and an empirical one.

From the statistical point of view, we estimated a unique supply function for the whole period. If it was misspecified, that is if the magnitude of the demand regime in 1994 was over exaggerated or even if the opposite regime should have occurred, the quality of the model would dramatically deteriorate in the period of time when the supply function is observable, that is in the supply regime. Obviously, this is not the case: the concordance between adjusted and historical values in the supply regime is as good as for the rest of the period.

From the empirical point of view, what needs to be stressed is that the level of credit, as measured by the ratio of credit to the corporate sector to GDP (see Table 7), was extremely low, hardly reaching 17.1 per cent of GDP in 1994. Therefore, a potential doubling of its volume was not unlikely to happen. Moreover, in the same year, there was an almost 40 per cent decline of the ratio as compared to 1989, and a 25 per cent as compared to 1991. It seems that banks wanted to again build up their market shares on a healthy basis, after the occurrence of a credit crunch in 1993.

Table 7. Credit to the Corporate Sector to GDP Ratio

1989	1990	1991	1992	1993	1994	1995	1996	1997	1998	1999
28.3	21.1	23	20.6	19.9	17.1	16.4	17.6	18.6	19.8	21.2

Source: National Bank of Poland and the authors' calculations.

5 Linking the Bank Lending Channel and the Disequilibrium Loan Market Analysis

We assumed in our theoretical model of the bank lending channel that the loan interest rate is perfectly flexible, thus clearing the loan market. However, the previous section showed the existence of a strong disequilibrium since 1994. Therefore, we should empirically investigate whether the interest rate spread is still a good indicator of the bank lending channel in the presence of a disequilibrium which, by definition, implies an imperfect adjustment of the loan interest rate. To this end, we applied

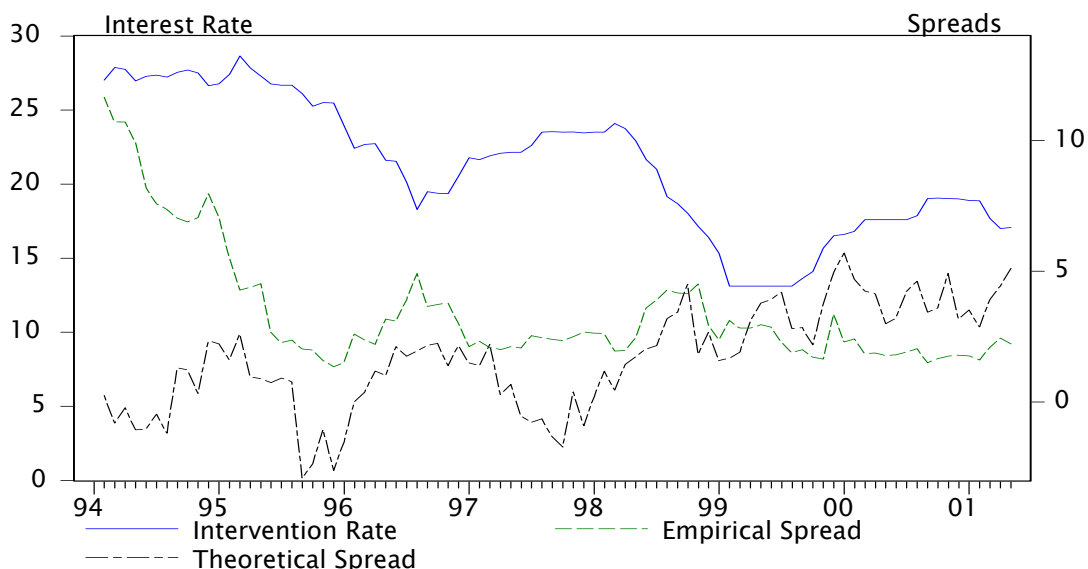
the following three-stage methodology.

First, using the model 3 estimates, we computed a theoretical price of loans that would have ensured an equality of the annual growth rates of loan demand and supply, other things being equal. Assuming that $\dot{D}_t = \dot{S}_t$ implies $D_t = S_t$, this theoretical loan interest rate is also the one which would have balanced the loan market in level.

Second, we constructed an interest rate spread between the theoretical loan rate and the empirical intervention rate.

Third, we compared the reaction of both the theoretical and empirical spreads to monetary policy actions reflected by the evolution of the empirical intervention rate.

Figure 7. Theoretical and Empirical Interest Rate Spreads and the Intervention Rate, II/94 - V/01



Source: National Bank of Poland and the authors' calculations.

As it appears in Figure 7, both spreads tend to change in the same direction, at least since the end of 1994. This indicates that in the presence of a loan market disequilibrium, the empirical interest rate spread is still a valuable indicator for distinguishing between amplification and attenuation effects of the bank lending channel on monetary policy shocks. However, the magnitude of its reaction is smaller as compared to the theoretical spread. Indeed, since January 1995 the variance of the latter amounts to 3.67 whereas it is equal to 1.09 for the former.

Conclusion

In this paper we have developed an extended AD/AS credit-augmented model of the transmission mechanism. The extensions include an interest rate control with flexible prices and an imperfect

nominal wage indexation. Using this framework, we established that the bank lending channel may amplify as well as attenuate the action of the traditional interest rate channel. We found that the change in the interest rate spread between a loan rate and a policy rate is a good indicator to gauge the efficiency of monetary policy impulses. If the pass-through of official interest rates to loan rates is less (more) than one, this implies a weaker (stronger) impact on prices and output: the bank lending channel will therefore reduce (increase) the potency of the traditional interest rate channel. An important explanatory factor of amplification and attenuation effects is the sensitivity of banks and firms to loan and market interest rates. But these results require two assumptions. First, there must be a positive relationship between loan rates and the policy rate, which is a highly probable outcome. Second, the loan interest rate should be a market clearing variable. We showed that although the latter assumption does not hold in the Polish case, the spread can still be used by the policymaker as a good indicator of attenuation and amplification effects of the bank lending channel. Our analysis indicates that the bank lending channel was reducing the overall potency of monetary policy from January 1996 to August 1998 and had and on average a neutral effect from September 1998 to June 2001.

Secondly, we have estimated a regime-switching model. The great number of tests we have performed makes us confident about the reliability of our results. The estimates reveal that disequilibrium is, in fact, a permanent characteristic of the Polish loan market since 1994. However, the type of regime changed over time. After a period of a strong and long-lasting demand regime until the end of 1997, the market definitely switched into a supply regime from December 1999.

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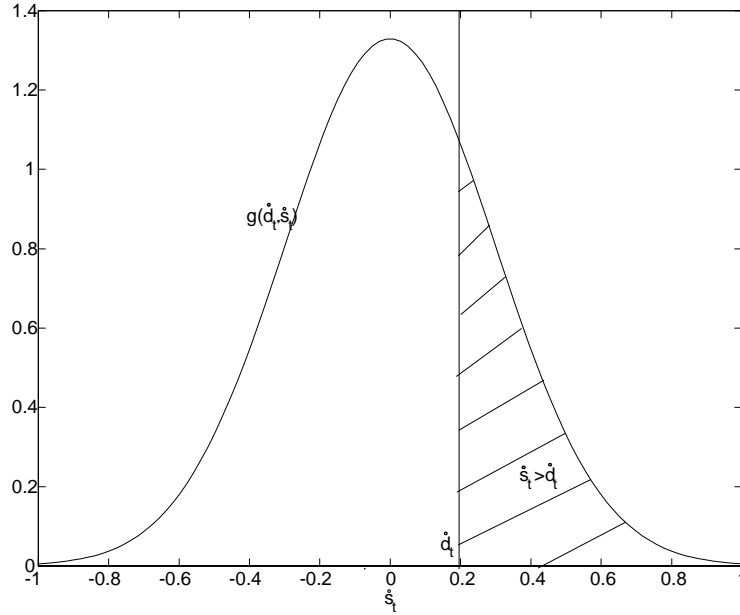
Appendix A. Marginal Densities of \dot{Q}_t

Let us denote $g_{\dot{D}_t, \dot{S}_t}(\dot{d}_t, \dot{s}_t)$ the joint density of \dot{D}_t and \dot{S}_t . We know that the corresponding marginal densities of the unobservable variables \dot{D}_t and \dot{S}_t are defined by:

$$f_{\dot{D}_t}(\dot{d}_t) = \int_{-\infty}^{\infty} g_{\dot{D}_t, \dot{S}_t}(\dot{d}_t, z) dz \quad f_{\dot{S}_t}(\dot{s}_t) = \int_{-\infty}^{\infty} g_{\dot{D}_t, \dot{S}_t}(z, \dot{s}_t) dz. \quad (\text{A-1})$$

We have to compute the marginal density of \dot{Q}_t on the two subsets $\dot{Q}_t = \dot{D}_t$, with $\dot{D}_t < \dot{S}_t$ and $\dot{Q}_t = \dot{S}_t$, with $\dot{S}_t < \dot{D}_t$. When $\dot{D}_t < \dot{S}_t$, for a given realization \dot{d}_t of \dot{D}_t , the marginal density of \dot{Q}_t is given by the area defined by the joint density $g_{\dot{D}_t, \dot{S}_t}(\dot{d}_t, z)$, for realizations \dot{s}_t superior to \dot{d}_t . The geometrical interpretation can be represented as on Figure 8.

Figure 8. Joint density $g_{\dot{D}, \dot{S}}(\dot{d}_t, \dot{s}_t)$ for a given \dot{d}_t



Under assumption that $\dot{D}_t < \dot{S}_t$, the marginal density of \dot{Q}_t is then given by:

$$f_{\dot{Q}_t | \dot{D}_t < \dot{S}_t}(\dot{q}_t) = \int_{\dot{q}_t = \dot{d}_t}^{\infty} g_{\dot{D}_t, \dot{S}_t}(\dot{d}_t, z) dz. \quad (\text{A-2})$$

Symmetrically, we get the marginal density of \dot{Q}_t when $\dot{S}_t < \dot{D}_t$:

$$f_{\dot{Q}_t | \dot{S}_t < \dot{D}_t}(\dot{q}_t) = \int_{\dot{q}_t = \dot{s}_t}^{\infty} g_{\dot{D}_t, \dot{S}_t}(z, \dot{s}_t) dz. \quad (\text{A-3})$$

Appendix B. Particular Case: $\sigma_{12} = 0$

In general case, we know that the marginal density of \dot{Q}_t is given by:

$$f_{\dot{Q}_t}(\dot{q}_t) = \int_{\dot{q}_t}^{\infty} g_{\dot{D}_t, \dot{S}_t}(\dot{q}_t, z) dz + \int_{\dot{q}_t}^{\infty} g_{\dot{D}_t, \dot{S}_t}(z, \dot{q}_t) dz, \quad (\text{B-1})$$

where the joint density $g_{\dot{D}_t, \dot{S}_t}(\dot{d}_t, \dot{s}_t)$, under assumption H_1 , corresponds to:

$$g_{\dot{D}_t, \dot{S}_t}(\dot{d}_t, \dot{s}_t) = \frac{1}{2\pi\sqrt{\det(\Omega)}} \exp \left[-\frac{1}{2} \begin{pmatrix} \dot{d}_t - x'_{1,t}\beta_1 \\ \dot{s}_t - x'_{2,t}\beta_2 \end{pmatrix}' \Omega^{-1} \begin{pmatrix} \dot{d}_t - x'_{1,t}\beta_1 \\ \dot{s}_t - x'_{2,t}\beta_2 \end{pmatrix} \right].$$

Let us assume that both residuals ε_1 and ε_2 are independent ($\sigma_{12} = 0$). In this case, the joint density can be expressed as the following simple expression:

$$\begin{aligned} g_{\dot{D}_t, \dot{S}_t}(\dot{d}_t, \dot{s}_t) &= \frac{1}{2\pi\sigma_1\sigma_2} \exp \left\{ -\frac{1}{2} \left[\left(\frac{\dot{d}_t - x'_{1,t}\beta_1}{\sigma_1} \right)^2 + \left(\frac{\dot{s}_t - x'_{2,t}\beta_2}{\sigma_2} \right)^2 \right] \right\} \\ &= \frac{1}{2\pi\sigma_1\sigma_2} \exp \left[-\frac{1}{2} \left(\frac{\dot{d}_t - x'_{1,t}\beta_1}{\sigma_1} \right)^2 \right] \times \exp \left[-\frac{1}{2} \left(\frac{\dot{s}_t - x'_{2,t}\beta_2}{\sigma_2} \right)^2 \right]. \end{aligned}$$

Now, consider the first member of the marginal density of \dot{Q}_t (equation B-1):

$$\begin{aligned} \int_{\dot{q}_t}^{\infty} g_{\dot{D}_t, \dot{S}_t}(\dot{q}_t, z) dz &= \frac{1}{2\pi\sigma_1\sigma_2} \int_{\dot{q}_t}^{\infty} \left\{ \exp \left[-\frac{1}{2} \left(\frac{\dot{q}_t - x'_{1,t}\beta_1}{\sigma_1} \right)^2 \right] \times \exp \left[-\frac{1}{2} \left(\frac{z - x'_{2,t}\beta_2}{\sigma_2} \right)^2 \right] \right\} dz \\ &= \frac{1}{\sqrt{2\pi}\sigma_1} \exp \left[-\frac{1}{2} \left(\frac{\dot{q}_t - x'_{1,t}\beta_1}{\sigma_1} \right)^2 \right] \times \frac{1}{\sqrt{2\pi}\sigma_2} \int_{\dot{q}_t}^{\infty} \exp \left[-\frac{1}{2} \left(\frac{z - x'_{2,t}\beta_2}{\sigma_2} \right)^2 \right] dz. \end{aligned}$$

In the first term of this expression, we recognize the value of the $N(0, 1)$ density function at the particular point $(\dot{q}_t - x'_{1,t}\beta_1) / \sigma_1$. Indeed:

$$\frac{1}{\sqrt{2\pi}} \exp \left[-\frac{1}{2} \left(\frac{\dot{q}_t - x'_{1,t}\beta_1}{\sigma_1} \right)^2 \right] = \phi \left(\frac{\dot{q}_t - x'_{1,t}\beta_1}{\sigma_1} \right),$$

where $\phi(\cdot)$ denotes the $N(0, 1)$ density function. Since this function is symmetric, the first member of the marginal density of \dot{Q}_t can be expressed as:

$$\int_{\dot{q}_t}^{\infty} g_{\dot{D}_t, \dot{S}_t}(\dot{q}_t, z) dz = \frac{1}{\sigma_1} \phi \left(\frac{x'_{1,t}\beta_1 - \dot{q}_t}{\sigma_1} \right) \times \frac{1}{\sqrt{2\pi}\sigma_2} \int_{\dot{q}_t}^{\infty} \exp \left[-\frac{1}{2} \left(\frac{z - x'_{2,t}\beta_2}{\sigma_2} \right)^2 \right] dz.$$

The second term of this expression can be transformed in order to introduce the $N(0, 1)$ cumulative distribution function, denoted $\Phi(\cdot)$. Indeed, let us consider the following change in variable $\tilde{z} = (z - x'_{2,t}\beta_2) / \sigma_2$, with $dz = d\tilde{z} \sigma_2$. Then, we have:

$$\begin{aligned} \frac{1}{\sqrt{2\pi}\sigma_2} \int_{\dot{q}_t}^{\infty} \exp \left[-\frac{1}{2} \left(\frac{z - x'_{2,t}\beta_2}{\sigma_2} \right)^2 \right] dz &= \frac{1}{\sqrt{2\pi}\sigma_2} \int_{\tilde{q}_t}^{\infty} \exp \left(-\frac{\tilde{z}^2}{2} \right) d\tilde{z} \sigma_2 \\ &= \frac{1}{\sqrt{2\pi}} \int_{\tilde{q}_t}^{\infty} \exp \left(-\frac{\tilde{z}^2}{2} \right) d\tilde{z}, \end{aligned}$$

with $\tilde{q}_t = (\dot{q}_t - x'_{2,t}\beta_2) / \sigma_2$. Then, this integral can be expressed as function $\Phi(\cdot)$:

$$\frac{1}{\sqrt{2\pi}\sigma_2} \int_{\dot{q}_t}^{\infty} \exp\left[-\frac{1}{2}\left(\frac{z - x'_{2,t}\beta_2}{\sigma_2}\right)^2\right] dz = 1 - \Phi(\tilde{q}_t) = \Phi(-\tilde{q}_t).$$

Finally, we get:

$$\int_{\dot{q}_t}^{\infty} g_{\dot{D}_t, \dot{S}_t}(\dot{q}_t, z) dz = \frac{1}{\sigma_1} \phi\left(\frac{x'_{1,t}\beta_1 - \dot{q}_t}{\sigma_1}\right) \Phi\left(\frac{x'_{2,t}\beta_2 - \dot{q}_t}{\sigma_2}\right).$$

Symmetrically, we can compute the second term of the marginal density of \dot{Q}_t (equation B-1) as:

$$\int_{\dot{q}_t}^{\infty} g_{\dot{D}_t, \dot{S}_t}(z, \dot{q}_t) dz = \frac{1}{\sigma_2} \phi\left(\frac{x'_{2,t}\beta_2 - \dot{q}_t}{\sigma_2}\right) \Phi\left(\frac{x'_{1,t}\beta_1 - \dot{q}_t}{\sigma_1}\right).$$

Then, the unconditional density of \dot{Q}_t is defined by equation (28):

$$\begin{aligned} f_{\dot{Q}_t}(\dot{q}_t) &= \frac{1}{\sigma_1} \phi\left(\frac{x'_{1,t}\beta_1 - \dot{q}_t}{\sigma_1}\right) \Phi\left(\frac{x'_{2,t}\beta_2 - \dot{q}_t}{\sigma_2}\right) \\ &\quad + \frac{1}{\sigma_2} \phi\left(\frac{x'_{2,t}\beta_2 - \dot{q}_t}{\sigma_2}\right) \Phi\left(\frac{x'_{1,t}\beta_1 - \dot{q}_t}{\sigma_1}\right). \end{aligned}$$

Appendix C. Data Description; an Alternative Specification of Model 3

- The loan rates are weighted averages for minimum interest rates applied to Polish firms by major banks. The sample of banks includes from 20 banks in February 1994 to 12 in June 2001 (the difference due to the consolidation process) and represents, on average for the February 1994 - June 2001 period, 75.2 per cent of the banking system in terms of total corporates' banking debt and 84.7 per cent in terms of total deposits.
- The intervention rate is a weighted average of 1 to 14-days reverse repo operation rates and that of the central bank securities issued for different maturities between February 1994 and January 1998 and is equal since then to the actual rate on 28-day NBP bills (also taking into account the average rate of outright operations).
- The industrial production is issued from the Monthly Bulletin publication of the Central Statistical Office (GUS).
- The loan series is for zloty denominated loans up to one year extended to resident and non resident firms for the above sample of banks.

Table 8. ML Estimates of an Alternative Specification of Model 3

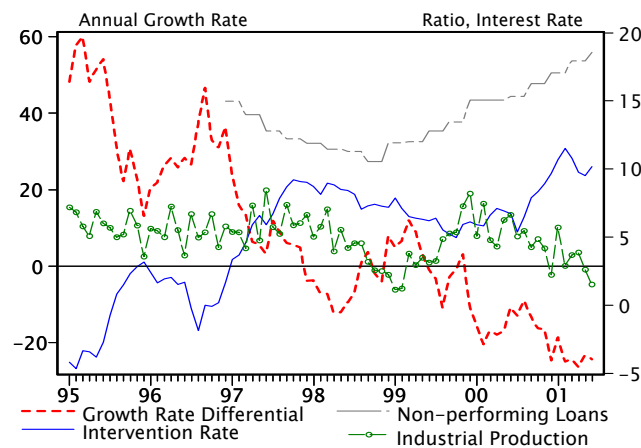
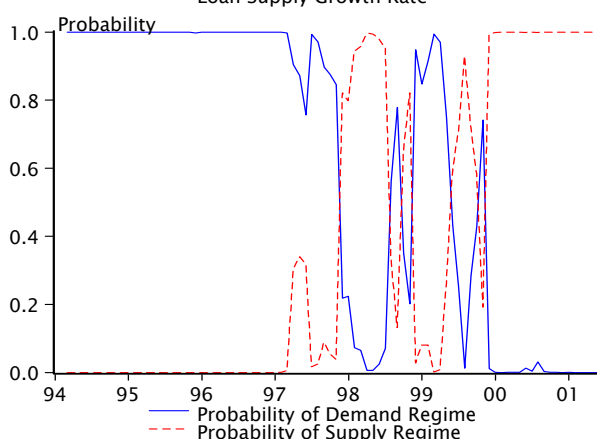
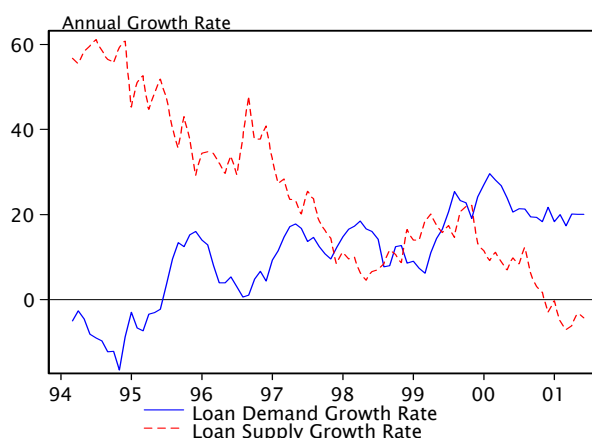
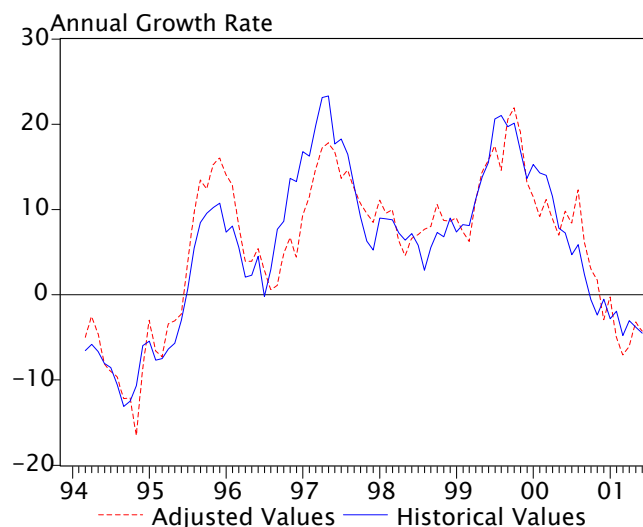
Variable	Demand Function	Supply Function
Constant	46.27 (9.05)	-37.92 (-2.68)
$(IL - IC)_{t-1}$	-0.78 (-3.00)	9.77 (3.47)
IC_{t-1}	-0.33 (-1.21)	0.02 (0.03)
$PROD_t$	0.20 (2.26)	—
DEP_t	—	1.85 (7.45)
$LGOR_t$	—	0.46 (6.60)
SAV_t	-0.28 (-6.15)	—
OL_t	-2.27 (-6.45)	—
σ_i^2	3.33	5.44
$R^2 = 0.80$	$\bar{R}^2 = 0.77$	$L = -243.45$

Asymptotic t-statistics are in parentheses, L : log-likelihood.

Appendix D. Model 3 with CPI Adjusted Variables

Variable	Demand Function	Supply Function
Constant	12.46 (5.53)	24.48 (4.41)
IL_{t-1}	-1.14 (-2.03)	1.57 (1.10)
IC_{t-1}	2.85 (6.42)	-5.47 (-4.20)
$PROD_t$	-0.03 (-0.25)	—
DEP_t	—	0.94 (4.87)
$LGOR_t$	—	0.12 (2.21)
SAV_t	-0.11 (-2.38)	—
OL_t	-2.16 (-4.93)	—
σ_i^2	3.72	3.05
$R^2 = 0.85$	$\bar{R}^2 = 0.83$	$L = -230.06$

Asymptotic t-statistics are in parentheses, L : log-likelihood.

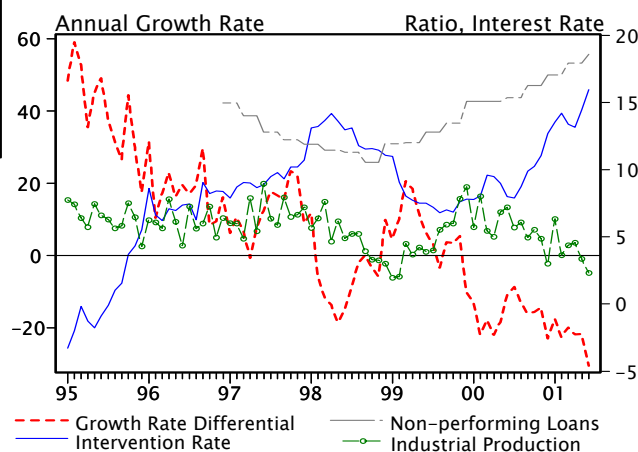
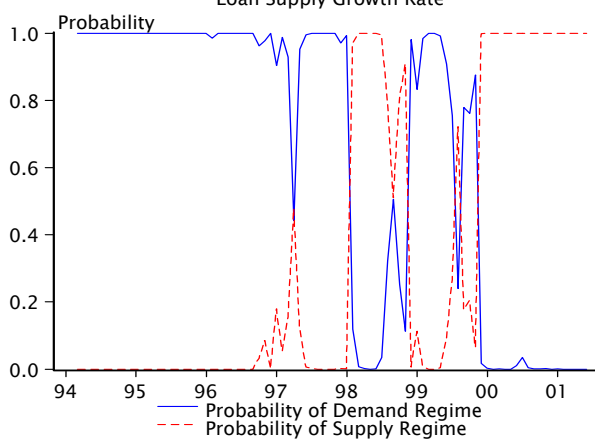
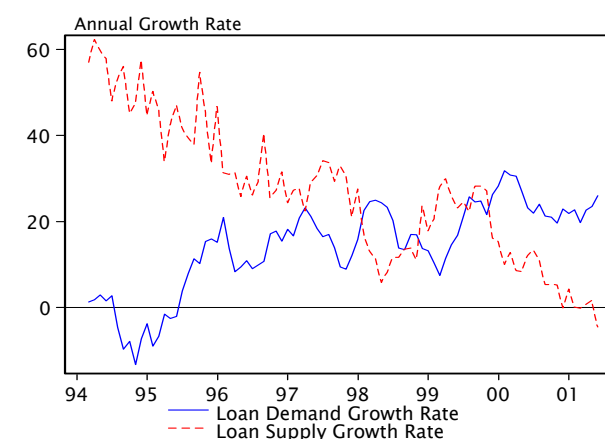
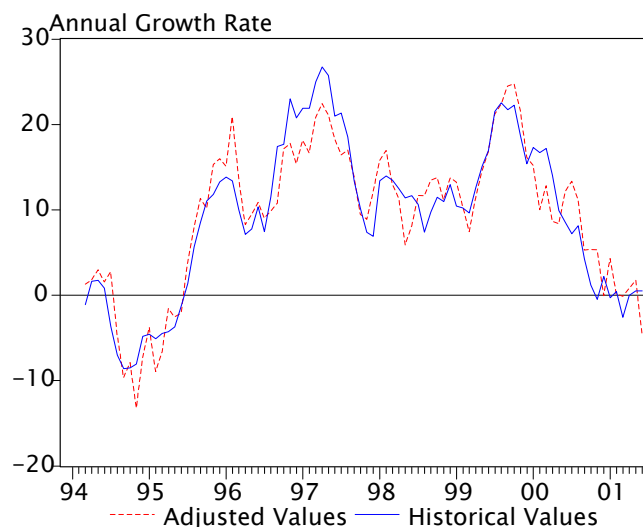


Source: National Bank of Poland, Central Statistical Office (GUS) and the authors' calculations.

Appendix E. Model 3 with PPI Adjusted Variables

Variable	Demand Function	Supply Function
Constant	9.46 (4.44)	16.40 (2.82)
IL_{t-1}	-1.30 (-4.20)	3.19 (2.00)
IC_{t-1}	3.16 (10.05)	-6.20 (-3.95)
$PROD_t$	0.007 (0.05)	—
DEP_t	—	1.52 (6.38)
$LGOR_t$	—	0.20 (3.68)
SAV_t	-0.10 (-2.37)	—
OL_t	-2.34 (-6.87)	—
σ_i^2	3.13	3.55
$R^2 = 0.87$	$\bar{R}^2 = 0.85$	$L = -226.98$

Asymptotic t-statistics are in parentheses, L : log-likelihood.



Source: National Bank of Poland, Central Statistical Office (GUS) and the authors' calculations.