Credit demand asymmetry in the Netherlands 1983-1997

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1 INTRODUCTION

In the past few years, there has been increasing attention for various kinds of asymmetries in the monetary transmission mechanism. Is the impact of monetary policy asymmetric over the business cycle? Does a monetary contraction have a stronger impact than a monetary expansion? Is the transmission of large interest rate changes different from that of small steps? Obviously, these issues are important to take into account in the formulation of monetary policy. In this paper, we investigate asymmetries in the credit demand equation. We do this by giving an overview of previous studies as well as presenting new results based on Dutch data. In line with most empirical work in this field (e.g. Fase, 1995; Kakes, 2000) we focus on the demand side of the credit market. This choice can be further motivated by empirical evidence that, at least in the short run, the Dutch credit market is demand determined (see e.g. Kakes, 2000) and can be characterised as a customer market (Swank, 1994).

Prior to the start of EMU, the Dutch credit market was a relevant factor for monetary policy, despite the fact that monetary autonomy was significantly reduced because of the fixed exchange rate *vis-à-vis* the Dmark and the high degree of capital mobility. Due to the imperfect integration of the credit market with international capital markets there was even scope for a domestic money supply target (see Kuipers and Sterken, 1995), which was actually pursued in some way or another until 1991. Is it still interesting to focus on the Dutch credit market after the start of EMU? After all, monetary policy is first and foremost based on euro-wide aggregated data now. However, credit markets of individual member states are still likely to exhibit specific national characteristics (see e.g. De Bondt, 2000). Hence, as cross-national differences in monetary transmission may complicate the implementation of a common monetary policy, it is still useful to consider individual EMU economies. Furthermore, credit growth is also interesting from a financial stability point of view, which has clearly a national dimension.

We investigate two sources of non-linearity in the credit demand equation: monetary policy asymmetry and business cycle asymmetry. Our results show that in periods of expansionary monetary policy, interest rate changes have more impact on aggregate credit than during monetary contractions. This result corroborates the findings of Van Ees et al. (1999) and Kakes (2000). Rather than making an a priori distinction between positive and negative monetary policy shocks, as these two studies do, we employ a threshold regression model and explicitly test whether nonlinearities are present. Interestingly, our conclusion is different from the standard 'credit view' literature. In particular, studies emphasizing the importance of a 'financial accelerator' (e.g. Bernanke et al., 1994) imply an asymmetry in the opposite direction. In line with Kakes (2000), we did not find evidence for business cycle asymmetry in the credit demand equation in the Netherlands.

This paper is organised as follows. In Section 2, we present an overview of previous literature on asymmetric effects of monetary policy. In Section 3, we briefly discuss the methodology we applied, and in Section 4 we present our results. Section 5 concludes.

2 NONLINEARITIES IN THE MONETARY TRANSMISSION MECHANISM

Is the monetary transmission mechanism characterised by asymmetric effects? This rather old theme in monetary economics goes back to pre-war theories. An early example of monetary transmission being dependent on the state of the economy is the liquidity trap, first described by Keynes (1936) and recently restated by Krugman (1998). Another example is that the impact of monetary policy may be different over the business cycle. Credit market imperfections and downward price rigidities may cause monetary policy to have more of an impact in a recession than during a boom (see Walsh, 1998 or Kakes, 2000, for a discussion of underlying theories). Furthermore, interest rate changes themselves may have asymmetric effects. It is often stated that especially monetary contractions have a real impact, whereas monetary expansions are ineffective (like 'pushing on a string'). Hence, monetary policy cannot be used to stimulate the economy or, as it was formulated in the 1930s, 'one can lead a horse to the water but one cannot make him drink' (Edie, 1931).

In recent years, many empirical studies have investigated whether these asymmetries can be observed in practice. Asymmetry over the business cycle has been established for the United States by Thoma (1994), Gertler and Gilchrist (1994), Kashyap et al. (1994), Garcia and Schaller (1995) and Weise (1999). Kakes (2000) finds asymmetric effects of monetary policy over the business cycle for the United States, Germany, the UK and Belgium, but not for the Netherlands. Asymmetric effects on real activity between monetary contractions and expansions have been established by Cover (1992) for the United States and by Karras (1996a, 1996b) and Kakes (2000) for various European countries and the United States.

Several studies have found that credit market imperfections play an important role in the monetary transmission process. These include analyses based on aggregate data (e.g. Bernanke and Blinder, 1992 and, for the Netherlands, Garretsen and Swank, 1998a, 1998b, and Kakes, 2000) as well as micro-data (e.g. Fazzari et al., 1988; Whited, 1992; Kashyap and Stein, 1999 and, for the Netherlands, Van Ees et al., 1998). Most of these studies do not explicitly consider asymmetric effects on credit aggregates, though, but focus on one or more of the underlying conditions (e.g. heterogeneity among borrowers). Only a few studies analyse credit market asymmetries in relation to the business cycle or monetary policy directly. Asea and Blomberg (1998) investigate how banks' loan standards change over the business cycle. They find that during the expansionary regime lending conditions are relaxed while during recessions they are tightened, which causes a pro-cyclical effect of credit supply. Interestingly, their results also imply that the impact of banks' lending policy on the economy is particularly great during booms, which is in contrast with the general view that credit market imperfections have a stronger impact in recessions (as follows from e.g. the 'financial accelerator' literature). Asea and Blomberg stress that their results imply the opposite of Akerlof's 'lemon' problem, according to which bad projects drive out the good ones. Rather, asymmetric information in credit markets may cause good projects to draw in bad projects during a boom. Azariadis and Smith (1998) develop a general equilibrium model which provides a theoretical foundation of Asea and Blomberg's results.

Van Ees et al. (1999) and Kakes (2000) find that expansionary monetary policy measures have more of an impact on credit growth than monetary contractions in the Netherlands. Van Ees et al. analyse the instalment and withdrawal of direct credit constraints, a policy instrument that was used by the Dutch central bank until 1991, and conclude that only the latter had a significant effect. Kakes (2000) finds the same result for a group of European countries (including the Netherlands) but not for the United States, using several monetary policy indicators. After a description of our methodology, we provide new evidence in the spirit of these two studies in Section 4. An important extension of our approach is that we do not make an a priori distinction between monetary expansions and contractions. Rather, we employ a threshold model in which the border between the expansionary and contractionary regime does not necessarily coincide with a distinction between interest rate increases and interest rate decreases, but may also be at a point slightly different from zero (e.g. increases higher than *x*% versus decreases and increases below *x*%).

3 THRESHOLD REGRESSION MODELS

Threshold regression models are a popular vehicle in non-linear modelling. They can be applied in models with multiple equilibria, or in situations where the sample must be split on the basis of a (continuously-distributed) variable. Recently, Bruce Hansen developed the statistical theory of threshold models further and made it accessible to a more general audience in a series of articles (Hansen, 1997, 1999a, 1999b and 2000). The exposition of threshold regression models in this section draws heavily on the Hansen papers.

Suppose we have n observations of the endogenous variable y_t, the m-vector of predetermined variables x_t and the threshold variable q_t. A threshold regression (TR) model takes the following form:

$$y_t = q_1' x_t + e_t \qquad\qquad q_t \le g \quad , \tag{1a}$$

$$y_t = q_2' x_t + e_t \qquad\qquad q_t > g \quad , \tag{1b}$$

where e_t is the regression error. The threshold variable q_t splits the sample in two regimes: the parameters may differ depending on the regime, which are determined by the threshold value ?.

The threshold regression model of Equations (1a)-(1b) can be expressed in a single equation

$$y_t = q_1' x_t I(q_t \le g) + q_2' x_t I(q_t > g) + e_t \tag{2}$$

where $I(.)$ denotes the indicator function, which has the value one if the argument is true and zero if the argument is false. Defining

$$x_t(g) = \left(x_t I(q_t \le g) \quad x_t I(q_t > g)\right)'$$

Equation (2) becomes

$$y_t = x_t(g)' q + e_t , \tag{3}$$

where $q = (q_1' \quad q_2')'$. The regression parameters in this equation, ? and ?, can be estimated by least squares (LS). Note that the LS estimator is equivalent to the maximum likelihood estimator (MLE) when the regression error e_t is iid $N(0, s^2)$. The easiest method to obtain the LS estimates is to use sequential conditional LS: compute the sum of squared residuals for all values of the threshold variable to find the minimum. More formally, for a given value of ? the LS estimate of ? is

$$\hat{q}(g) = \left(\sum_{t=1}^{n} x_t(g) x_t(g)'\right)^{-1} \left(\sum_{t=1}^{n} x_t(g) y_t\right), \tag{4}$$

with residuals $\hat{e}_t = y_t - x_t(g)'\hat{q}(g)$ and variance

$$\hat{s}_n^2(g) = \frac{1}{n}\sum_{t=1}^{n}\hat{e}_t(g)^2. \tag{5}$$

The LS estimate of ? is the value that minimises the sum of squared errors of Equation (6)

$$\hat{g} = \arg\min_{g \in \Gamma} \hat{s}_n^2(g), \tag{6}$$

for $g \in \Gamma = \left[\underline{g}, \bar{g}\right]$, where $\underline{g}$ and $\bar{g}$ are the lower and upper value of the threshold variable. If the number of observations, n, is very large, G can be approximated by a grid. The LS estimate of *?* is computed as $\hat{q} = \hat{q}(\hat{g})$.

An important question is whether the threshold regression model of Equations (1a-1b) is statistically significant to the linear alternative, which has the null hypothesis H_0: $?_1 = ?_2$. In this situation the threshold parameter is not identified under the null hypothesis, which makes the testing problem complex. Hansen (1996) proposes a (heteroskedasticity-consistent) Lagrange Multiplier test. The p-values are computed by a bootstrap analogue, fixing the regressors from the right-hand side of the equation that is estimated and generating the bootstrap dependent variable from the distribution $N\left(0, \hat{e}_t^2\right)$, where $\hat{e}_t$ is the OLS residual from the estimated threshold regression model. Hansen (1996) shows that this bootstrap analogue produces asymptotically correct p-values.

Confidence intervals

Valid confidence intervals for the threshold parameter can be based on the likelihood ratio (or F) statistic

$$LR_n(g) = n\frac{\hat{s}_n^2(g) - \hat{s}_n^2(\hat{g})}{\hat{s}_n^2(\hat{g})}, \tag{7}$$

which tests the null hypotheses $H_0 : g = \hat{g}$. The null hypothesis is rejected for large values of $LR_n(g_0)$. Note that the likelihood ratio statistic is equal to zero at $g = \hat{g}$. Confidence intervals for the threshold parameter are constructed by inverting the likelihood ratio statistic $LR_n(g)$. The asymptotic critical values for this statistic are listed in Hansen (1997, Table 1) and Hansen (2000, Table 1). A graphical method to find the confidence interval for the threshold parameter is to plot the likelihood ratio statistic $LR_n(g)$ for all values of *?* and to check for which values of *?* the statistic crosses the horizon line that belongs to the confidence level of the test. The confidence region is constructed under the assumption that the error terms are homoskedastic. If the conditional homoskedasticy condition $E\left(e_t^2 \mid q_t\right) = s^{\ 2}$ does not hold, the same analysis can be done with an appropriately scaled likelihood ratio statistic to which an amended confidence region belongs.

The distribution of the parameters $\hat{q} = \left(\hat{q}\left(\hat{g}\right)\right)$ can be approximated by the conventional normal approximation if the threshold value ? is known with certainty. However, the uncertainty regarding the estimation of ? needs to be taken into account, especially in small samples like ours. Hansen proposes to do this as follows. Construct a ß-level confidence interval for ?, calculate the confidence interval for ? for all ?'s in this interval, and take these intervals together.

4 RESULTS

We construct a threshold regression equation for credit in the Netherlands, which we interpret as representing the demand side of the credit market. This is consistent with the signs we find for the interest rate coefficients (insofar as these are significant), but also with studies that conclude that the credit market is – at least in the short run – demand determined (Kakes, 2000) and shows the characteristics of a customer market (Swank, 1994). We use quarterly observations from the first quarter of 1983 up to and including the final quarter of 1997. This is a homogeneous period for monetary policy, given the fact that the Dutch guilder was pegged to the Dmark at a constant parity. Furthermore, it is important to note that we do not include the late 1990s, which has been a period of very high credit growth in combination with decreasing interest rates. Hence, our results cannot be attributed to this specific period. Typically, credit demand (*CR*) depends on output (*GDP*), a long-term interest rate (r_l) and a short-term interest rate (r_s).[3]

Two general remarks can be made about the role of interest rates in our analysis. First, because the short-term interest rate is largely controlled by the central bank, being its main operational target, this variable is often considered an accurate measure of monetary policy stance. Note, however, that the short-term interest rate also reflects endogenous responses of the central bank to

[3] We use the following data and sources. Credit is domestic credit, published in the IMF's International Financial Statistics (IFS), which includes credit by banks and institutional investors. Output is GDP published by Statistics Netherlands. The long-term interest rate is the government bond yield and the short-term interest rate is the overnight interest rate, both taken from the IFS. Real credit and real GDP are obtained using the GDP deflator, also published by Statistics Netherlands. All variables, except interest rates, are taken in natural logarithms.

developments in the economy. One way to circumvent this endogeneity problem would be to use VAR-based measures to isolate the unanticipated part of monetary policy.[4] This would be difficult to implement in our single-equation setting, though. Furthermore, Kakes (2000) shows that, for the Netherlands, monetary policy measured by the short-term interest level or by VAR-based interest rate shocks have a similar impact on credit and other variables. Our second remark is that the interest rates we include are only indirectly related to the lending rates that actually play a role in the credit market. Lending rates are not directly observable and are not the same for each client due to differences in creditworthiness and other customer idiosyncracies.

The estimated equation looks as follows

$$y_t = q_1' x_t I(q_t \leq g) + q_2' x_t I(q_t > g) + e_t \,,$$

where the endogenous variable is credit and the set of regressors consists of a constant, lags of credit, (lags of) output and (lags of) the interest rate(s). Because we allow all parameters in the credit demand equation to be regime-dependent, while we believe it is unrealistic to assume that credit can be explained by either output or interest rates alone, the number of parameters to be estimated is large. As a consequence, we cannot include many lags. To keep the analysis tractable in terms of degrees of freedom, we set the maximum lag of the right-hand-side variables at two quarters.

We consider two threshold variables: (i) changes in the short-term interest rate, which can be interpreted as monetary policy, and (ii) business cycle changes. For these threshold variables we investigated several specifications distinguishing: (i) nominal or real credit and GDP; (ii) variables in levels, first differences and annual differences; and (iii) the threshold variable in first differences or in annual differences. The threshold variable was allowed to have a lag between 0 and 6 periods. For all specifications we carried out the linearity test described in Section 3, using 1000 replications to compute the bootstrap-calculated *p*-values. In the computations the top and bottom 25% of the sample of observations on the threshold variable were trimmed. There was no evidence of residual heteroskedasticity in any of the models we estimated, so we did not make a correction for this.

[4] Further refinements of this approach are structural VARs, which are more explicitly based on an a priori economic structure and take into account (changes in) the central bank's operating procedures over time (see e.g. Walsh, 1998, for a discussion). For instance, Bernanke and Mihov (1995) develop a structural VAR in which they model the federal funds market.

For all specifications where linearity is rejected, we computed the estimates of the parameters of the linear model and the threshold value and the parameters in both regimes. A general observation is that interest rates are hardly significant in the linear versions of the credit demand equation, in which credit is typically explained by the constant term, lagged credit and economic activity. Previous studies on the Dutch credit market have also found very low or insignificant interest rate elasticities, particularly for short-term credit (see Fase, 1990, for an overview).[5] However, interest rates enter the scene when we split the sample and estimate a threshold regression.

Table 1 presents the outcomes of the best threshold regression with the short-term interest rate as threshold variable.[6] Here, credit and GDP are in real terms, all variables are in levels except the threshold variable for which the annual difference is taken (with a three-quarter lag). The linearity null hypothesis is rejected with a *p*-value of 0.025. The LS estimate of the threshold value equals -0.30, with a 95% confidence interval of [-0.48; -0.08] as can be verified in Figure 1. Looking at the results for the two regimes, both interest coefficients are significant in regime 1 (annual interest short-term rate changes below the threshold value of -0.30%) and substantially higher in absolute terms than in the linear equation. By contrast, interest rates are insignificant in regime 2 (interest changes higher than -0.30%). Hence, the impact of both short-term and long-term interest rate changes on credit is substantial in a regime of declining short-term interest rates, whereas it is insignificant in a regime of only modestly declining or increasing short-term rates.

[5] Fase (1995) finds significant interest rate elasticities in a linear framework, but his approach differs from ours in several respects. He focuses on short-term credit whereas we consider aggregate credit; he estimates over a different sample (1970-1990) and his elasticity refers to the lending rate, whereas we include the short-term interbank rate and the government bond yield.

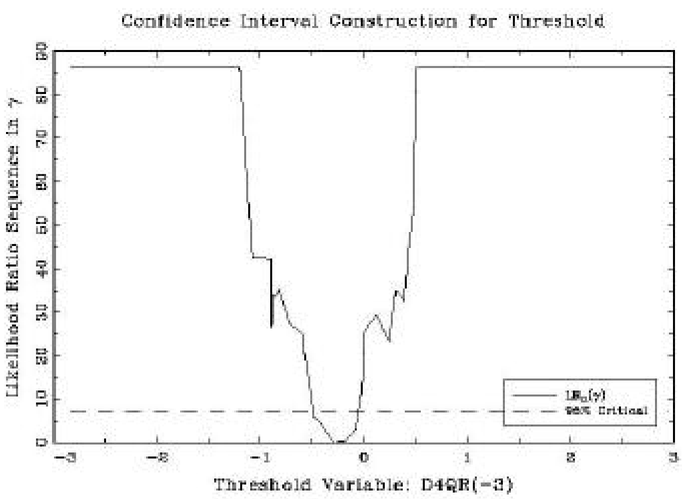
[6] This ‘best’ regression is the result of on several criteria. First, the *p*-value is required to be lower than 0.05. Second, the threshold value must be plausible. Third, there should be sufficient observations in both regimes.

Table 1 Threshold regression for real credit (*cr*):

threshold variable q_t =? $_4r_{s,t-3}$

	Linear model			Threshold regression		
			Regime I: $q_t \leq$ -.30		Regime II: q_t>-.30	
Variable	Estimate	St. error	Estimate	95% interval	Estimate	95% interval
Constant	-0.593	.293	-1.233	[-2.126; -.358]	-2.908	[-5.264; .017]
cr_{t-1}	.795	.164	.296	[-.078; .671]	.622	[-.084; 1.491]
cr_{t-2}	-.032	.165	-.029	[-.430; .400]	-.268	[-.898; .330]
gdp_t	0.805	.405	-.444	[-.541; 1.428]	1.187	[.030; 2.188]
gdp_{t-1}	-.458	.541	.597	[-.598; 1.890]	-.460	[-1.991; .873]
gdp_{t-2}	.037	.412	.038	[-.924; .919]	.611	[-.829; 1.757]
$r_{l,t}$	-.010	.007	-.00001	[-.027 .024]	-.005	[-.029; .145]
$r_{l,t-1}$	.002	.012	-.041	[-.037; -.005]	.013	[-.014; .046]
$r_{l,t-2}$	.004	.008	.015	[-.007; .036]	-.009	[-.037; .013]
$r_{s,t}$	-0.001	.006	-.013	[-.027; .004]	-.006	[-.028; .020]
$r_{s,t-1}$	.006	.010	.022	[-.003; .046]	-.003	[-.033; .022]
$r_{s,t-2}$	-.009	.006	-.027	[-.042; -.012]	-.013	[-.035; .016]
# of obs.	50		26		24	
SSE	.008		.0004		0.0012	
R^2	.993		.999		.996	

Figure 1 Confidence interval for threshold variable ? $_4r_{s,t-3}$



This is confirmed by Figure 2, in which the impact of 1%-point increases in the short-term and the long-term interest rates is simulated over ten quarters, using the estimated threshold regressions of both regimes in Table 2.[7] Presented are both the effect of a one-period increase, denoted as a 'temporary shock', and a permanent increase of the interest rates, denoted as a 'permanent shock'. Since these simulations also take into account the lag structure of the other variables, they give a more complete picture of the impact of interest rate changes than just the estimated interest rate coefficients.

Figure 2a Simulated impact short-term interest rate shock on credit

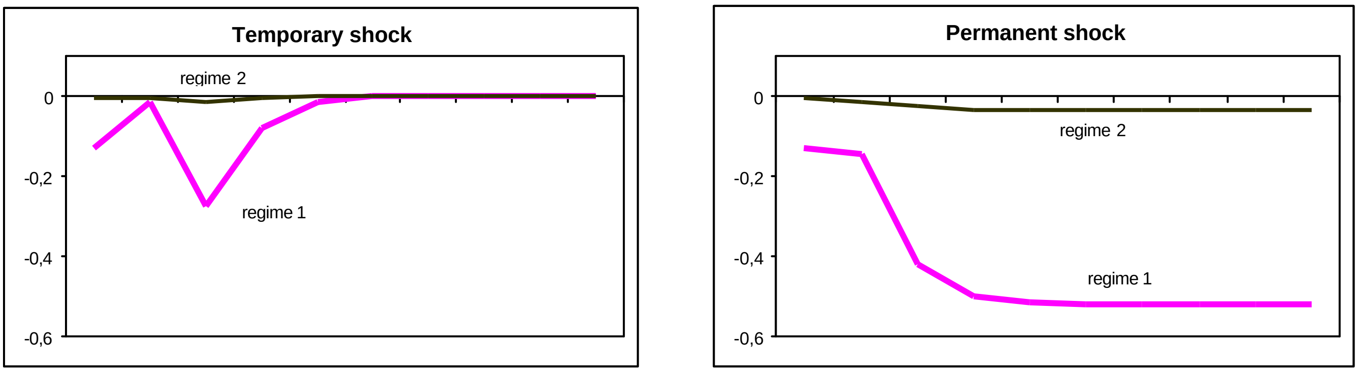
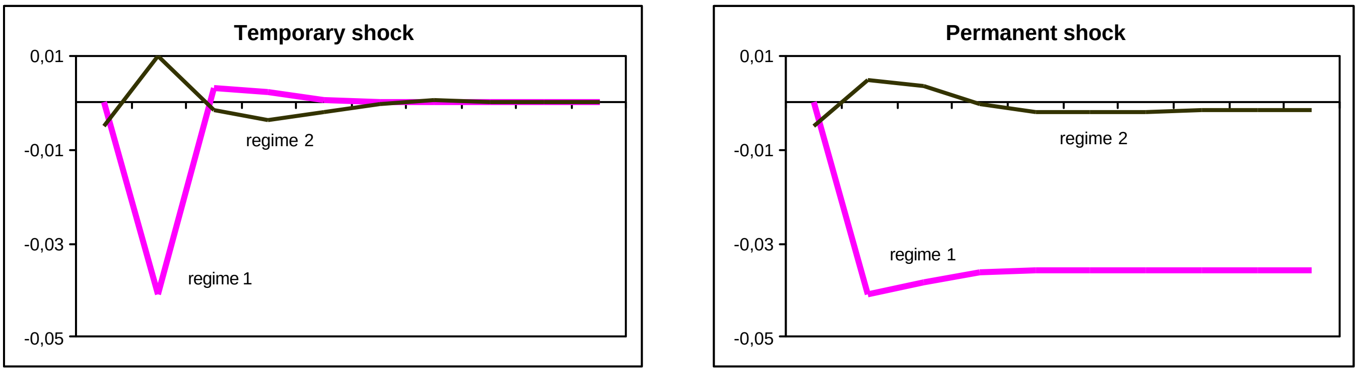


Figure 2b Simulated impact long-term interest rate shock on credit



Our result corroborates the conclusion of Van Ees et al. (1999) and Kakes (2000), that monetary policy has more impact on credit aggregates if monetary conditions are expansionary. This is also in line with the experience in the past few years, when credit growth increased markedly while

[7] Note that in all cases we simulate interest rate increases. Given the fact that the individual equations of each regime are linear, the effect of an interest rate decrease mirrors the simulation results that are

interest rates decreased. It is hard to explain this acceleration in credit growth with a linear model.[8]

How can this asymmetry be explained? Our findings are inconsistent with the standard credit view literature which implies the existence of asymmetric effects in the other direction, namely that a monetary contraction has more impact on credit than a monetary expansion. However, an explanation based on credit market imperfections along the lines of Asea and Blomberg (1998) may be useful to explain our results, by taking into account heterogeneity among borrowers. Following a monetary contraction, there may be a shift from lending to low-quality borrowers towards high-quality borrowers: the first are rationed, while the latter try to create a buffer of external funds, possibly anticipating a further contraction. At the aggregate level, these two effects may cancel out, which implies that total credit is not significantly affected by a monetary tightening. By contrast, following a monetary expansion loans are given more easily to both low-quality and high-quality borrowers (in the terminology of Asea and Blomberg, 'good projects draw in bad projects'), so in this case there will be a significant impact on aggregate credit. Unfortunately, it is difficult to test whether this explanation is relevant for the Netherlands, as disaggregated time series for several types of borrowers are not available.[9]

Another possible explanation for the observed asymmetry is that the Dutch credit market may be regarded as a 'customer market', rather than an auction market (Okun, 1981). An important characteristic of customer markets is that banks attach value to long-term client relationships, which implies that they shield their loans portfolio after a monetary contraction. An interview study among banks by Swank (1994) provides empirical support for the concept of a customer market in the Netherlands.

Detailed analysis of the regression outcomes with the business cycle as threshold variable showed that here the evidence of threshold effects is weak. In most of the cases the, grid search produced

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presented here. In addition, it should be noted that each simulation is based on only one equation, and thus does not take into account the possibility of regime switches during the simulation period.

[8] For example, the Dutch central bank's macroeconometric model MORKMON can only explain 60% of the increase in bank lending to the corporate sector in 1998 (see De Nederlandsche Bank, 1999).

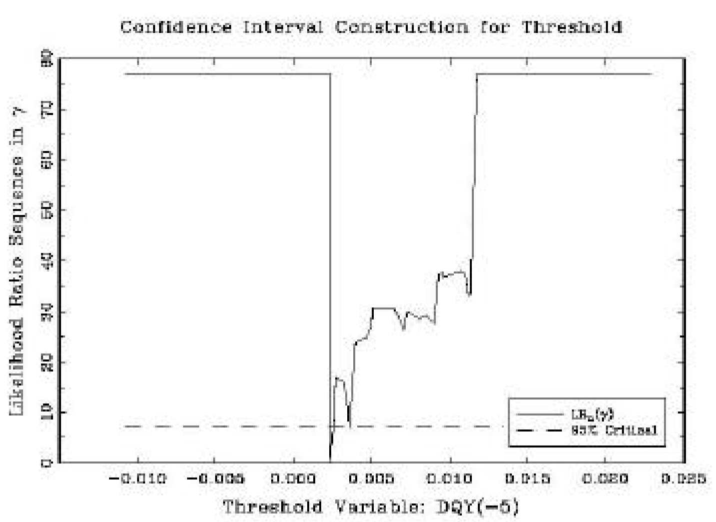
[9] The only distinction that can be made is between households and firms (see e.g. Garretsen and Swank, 1998b; Kakes, 2000). However, the interpretation of this distinction is not straightforward. On the one hand, the information asymmetry is larger for household credit than for corporate credit. Firms are generally better screened, particularly regarding expected cash flows. On the other hand, the bulk of loans to households consist of mortgages, which significantly reduces their risk. This would imply that households are for that part to be considered the 'high quality' borrowers.

corner solutions, which implies that the number of observations in one regime becomes so low that unrealistic parameter estimates are obtained. Table 2 lists the outcome with the lowest *p*-value: the combination of nominal credit and GDP, variables in levels and the first difference ? gdp_{t-5} as threshold variable. The LS estimate of the threshold value is 0.002 with 95% confidence interval [0.002; 0.003]. Figure 3 illustrates that this is a corner solution. Hence, we cannot conclude that there are asymmetries over the business cycle. This corresponds to the conclusion of Kakes (2000) who finds that business cycle asymmetry can be established for the United States, Germany, Belgium and the United Kingdom, but not for the Netherlands, using a Markov switching model to separate recessions and expansions.

Table 2 Threshold regression for nominal credit (*cr*):
threshold variable q_t =? gdp_{t-5}

	Linear model			Threshold regression		
			Regime I: $q_t \leq .002$		Regime II: q_t>.002	
Variable	Estimate	St. error	Estimate	95% interval	Estimate	95% interval
Constant	-0.244	.162	-.776	[-1.815; .262]	-.274	[-.558; .009]
cr_{t-1}	.699	.159	.086	[-.665; .836]	1.004	[.679; 1.329]
cr_{t-2}	.230	.159	.656	[-.149; 1.461]	-.117	[-.449; .215]
gdp_t	-0.076	.255	-.681	[-2.178; .816]	.224	[-.238; .686]
gdp_{t-1}	.119	.403	1.351	[-1.338; 4.041]	-.113	[-.787; .561]
gdp_{t-2}	.076	.261	-.269	[-1.800; 1.261]	.056	[-.411; .522]
$r_{l,t}$	-.002	.006	.035	[-.037; .107]	-.018	[-.029; -.007]
$r_{l,t-1}$	-.005	.009	-.054	[-.143; .034]	.020	[.003; .038]
$r_{l,t-2}$	.006	.006	.018	[-.012; .048]	-.010	[-.021; .001]
$r_{s,t}$	-0.0002	.005	.002	[-.023; .027]	.006	[-.004; .016]
$r_{s,t-1}$	.004	.008	.012	[-.041; .065]	-.002	[-.018; .014]
$r_{s,t-2}$	-.007	.005	-.025	[-.060; .010]	-.003	[-.012; .006]
# of obs.	53		14		39	
SSE	.0056		.0004		0.0019	
R^2	.998		.999		.999	

Figure 3 Confidence interval for threshold variable ? gdp_{t-5}



5 CONCLUDING REMARKS

In this paper we investigated whether a non-linear relationship between credit, output and interest rates in the Netherlands can be established for the 1983-1997 period. We found evidence of asymmetric effects of monetary policy on aggregate credit: if the monetary stance is expansionary, the impact of interest rate changes on credit is larger than in periods when the monetary stance is tightening. This finding is in line with previous studies based on different methodologies and different sample periods, and with the credit boom in the second half of the 1990s which coincided with decreasing interest rates. We could not establish asymmetry over the business cycle in the Netherlands for the 1983-1997 period. This finding is also consistent with previous research.

Two further remarks can be made. First, we have focused on the first stage of the monetary transmission process. We did not analyse the consequences for inflation or real activity. Nevertheless, credit growth is generally considered an important information variable for monetary policy. Our results thus imply that non-linearity in credit demand should be taken into account for the assessment of the monetary policy stance. Perhaps this would explain the finding of Garretsen and Swank (1998a), based on a linear model, that credit does not perform very well as a leading indicator for inflation and real activity in the Netherlands. Second, we did not explicitly investigate the underlying cause of the observed asymmetry. This would probably require an analysis with disaggregated data, in order to investigate the behaviour of different types of borrowers and lenders. The fact that our result differs from the standard 'credit view' is consistent with recent studies (Garretsen and Swank, 1998a, 1998b; Kakes, 1998b) which conclude that a bank lending channel is not important for the Netherlands.

In this paper we looked for asymmetries in credit demand in the Netherlands using a single-equation threshold regression model. The research can be extended in several directions. A next step may be to investigate asymmetries in credit demand using aggregated euro area data. In addition, other methods can be applied. Candidates are the smooth transition regression models advocated by Teräsvirta (1994, 1998), which assume a gradual transition from one regime into the other, or a multivariate threshold regression model (Tsay, 1998; Choi, 1999; Chen and Shiang, 1999; Weise, 1999).

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