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AN ARDL APPROACH TO IDENTIFY BANK LENDING CHANNEL IN INDONESIA

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Abstract

This paper tests whether the bank lending channel works in Indonesia. It develops an error correction representation of the Autoregressive Distributed Lag (ARDL) model of two bank credit markets. Each model takes account of one structural break associated with the 1998 financial crisis. The date of the crisis is determined by a unit root test that includes two structural breaks. Instead of Johansen's cointegrating procedure, bounds test procedure is implemented. The estimated error correction model for both markets suggests that bank loans adjust more strongly towards loan supply, implying that monetary-induced disturbances in bank loans originate from the supply side.

Keywords: *bank lending channel, unit root, structural breaks*
JEL classification numbers: *C12, C22, C52, F13*

INTRODUCTION

One of the channels through which a monetary policy change is transmitted to the level of economic activities and inflation is bank lending channel[1]. Bank lending channel arises from two basic assumptions (Gertler and Gilchrist, 1993). *First*, the central bank, through a change in monetary policy, is able to constrain the supply of bank loans. That is, a monetary contraction will reduce bank reserve money and in turn the quantity of loans. This implies a monetary-policy-induced decline in bank loans originates from the supply side, instead of the demand side. *Second*, some firms, mainly small ones, become dependent on bank loans because they find it prohibitively expensive to obtain funds from other means, such as issuing securities or bonds, due to high screening and monitoring costs. Accordingly, if banks become unable or unwilling to extend credit, there will be a fall in spending by dependent customers, and therefore, aggregate demand. Taken together, these two assumptions imply that a monetary policy contraction will reduce the supply of bank loans and in turn affect real economic activity.

Most empirical studies on bank lending channel that employ vector autoregression (VAR) models with either aggregate or disaggregated data provide ambiguous results[2]. Following a monetary policy tightening bank loans respond with a lag and decline almost contemporaneously with the aggregate output. These results fail to satisfy the first assumption of bank lending channel. They could not identify whether the monetary-policy-induced decline in bank loans originates from the supply side or demand side. Worst still, the fact that bank

[1] Other channels include interest rate, asset price, exchange rate and expected inflation channels. For a detailed exposition of each channel, see Mishkin (1995; 2001).

[2] See Bernanke (1986), Bernanke and Blinder (1988); Bernanke and Blinder (1992); Kashyap, Stein and Wilcox (1993), Ramey (1993), and Olliner and Rudebnusch (1996), to name but few.

loans decline almost contemporaneously with the decline in the aggregate output following a monetary policy tightening may indicate that the monetary-policy-induced decline in bank loans originates from the demand side. This means monetary policy works through interest rate channel, instead of bank lending channel. That is, a monetary policy tightening reduces money supply and raises interest rates thereby depressing economic activities, which in turn reduce the demand for credit. Therefore, there is an identification problem: the monetary-policy-induced movement in the bank loans is not identified as to whether it is demand or supply determined.

Kakes (2000) attempted to solve this identification problem by employing a five-variable vector error correction (VEC) model of bank credit market in the Netherlands. He assumed three cointegrating vectors identified as bank lending demand relation, bank lending supply relation and banks' bond holding relation. Identifying the first two cointegrating vectors as bank lending and supply relations can help solve the identification problem. That is, whether the credit market originates from the supply or demand sides depends on the short-run adjustment toward the equilibrium in the bank credit market in the VEC model. The short-run adjustment toward the credit market equilibrium is said to be dominated by supply of credit if in the equation of bank lending change the short-run adjustment coefficient on the error correction term corresponding to the long-run supply relation is greater in magnitude or statistically more significant than that on the error correction term corresponding to the long-run demand relation.

Agung et al. (2002) applied a similar approach to identifying bank lending channel in Indonesia. They develop a four-variable VECM that assumes two cointegrating vectors identified as bank lending supply and demand relations. While the bank lending demand relation is of the same specification as the one specified by Kakes (2000), the bank lending supply is positively related to the level of economic activity and the interest rate differential, which they assume to be the spread between the banks' lending rate and banks' funding costs proxied by deposit rate. Successfully identifying the cointegrating vectors, the former concluded that the credit market in the Netherlands is demand-determined, while the latter found that the credit market in Indonesia is supply-induced.

However, Agung *et al*. (2002) ignored the 1998 financial crisis, which was largely responsible for the credit crunch in the aftermath of the crisis. The existence of credit crunch might have reinforced the decline in the supply of bank loans following a monetary tightening. The credit crunch is said to be present only if the decline in the bank loans following the crisis is supply side phenomenon. This may be identified by including the shift dummy in the supply of credit relation only and expecting its coefficient to be negatively signed and significant. Alternatively, the shift dummy may enter both supply and demand relations and its estimated coefficient in the former is greater in magnitude or more significant than in the latter relation.

Another important drawback of those studies is that some variables of the Indonesian bank credit markets follow a stationary process. This is equally the case when a conventional Augmented Dickey-Fuller (ADF) unit root test or a unit root test that allows for one or more structural breaks is implemented. Since the variables in the system are not of the same degree of integration a conventional Johansen's VEC approach is not applicable. Likewise, another cointegration test procedure and modelling framework that accommodate a mixed set of variables is needed.

This study is intended to offer a solution to those problems in identifying bank lending channel of monetary policy transmission in Indonesia. In so doing, it implements both conventional ADF and new unit root tests to determine the degree of integration of the variables under study. Since the results of both unit root tests suggest that the variables are of mixed degree of integration, a new cointegration test procedure introduced by Pessaran et al. (2001) is implemented. Two long-run equations that represent bank lending supply and demand relations are assumed. For each equation a separate bounds test is implemented to determine whether the variables in each equation are cointegrated. The implementation of bounds test takes account of one structural break associated with the 1998 financial crisis. The exact date of the crisis will be endogenously determined by implementing a unit root test that allows for one structural brake. Further, an error correction representation is estimated to produce both estimated long-run bank lending supply and demand equations and the short-run adjustment mechanism in the system.The remainder of this paper is organized as follows. The following section discusses the empirical framework that includes the construction of a long-run model of the Indonesian bank credit markets, unit root test that allows for two structural breaks, bounds test procedure and data description. Section 3 presents the results and discussion. Finally section 4 concludes.

METHODS

ARDL-Based Error Correction Model

This study employs an error correction model, which is based on the autoregressive distributive lag (ARDL) framework introduced by Pessaran et al. (2001). Similar to Angle-Granger's approach, the ARDL-based error correction model also consists of two components: a long-run equilibrium equation and a short-run adjustment mechanism to the long-run equilibrium. Two error correction models are devised to capture the bank lending supply and demand in Indonesian bank credit market. The long-run bank lending supply is a function of the spread between the banks' lending rate and banks' lending opportunity costs proxied by call-money rate, and the level of economic activity. Another variable expected to influence the bank lending supply is the financial crisis proxied by a shift dummy. The use of call-money rate as a proxy for bank lending opportunity costs is due to the fact that for many commercial banks it serves as one of important sources of funds as well as a place to invest their excess supply of funds. It is expected that both spread and level of economic activity are positively related to the bank lending supply. Further, the long-run bank lending demand is a function of the level of economic activity and bank lending rate. While the former explanatory variable is expected to be positively related, the latter is expected to be negatively related, to the bank loans demand.

Since there are two bank credit markets in Indonesia: the working capital credit market and the investment credit market, two separate models are devised. Each model consists of five variables, namely the Real GDP (Y_t) as measure of the level of economic activity, bank loans, call-money rate ($R1_t$), lending rate, and the shift dummy (*SD9903*) that has zero for observations before the third month of 1999 and one thereafter[3]. While in the former market bank

[3] A unit root test that allows for two unknown break dates finds that these two markets share one common break date, the third month of 1999, which coincides with the period of the crisis. Inspection of the data also confirms this break date. These two markets differ in the other estimated break dates which fall before the crisis unfolded. The financial crisis is believed to start unfolding in August 1997 when a sharp depreciation of the Rupiah occurred. It did not turn

loans are represented by Working Capital Loans in real terms ($LRCWCRP_t$) and the lending rate by Working Capital Lending Rate (RWC_t), in the latter they are represented by Investment Capital Loans in real terms ($LRCINVRP_t$) and Investment Capital Lending rate ($RINV_t$), respectively. The specification of these markets is as follows.

Bank Working Capital Credit Market

Supply
$$LRCWCRP_t^S = a_0 + a_1 LY_t + a_2(RCW_t - R1_t) + a_3 SD9903 \quad (1)$$

Demand
$$LRCWCRP_t^D = b_1 LY_t + b_2 RWC_t \quad (2)$$

where it is expected that $a_1>0$, $a_2>0$, $a_3<0$, $b_1>0$, and $b_2<0$.

Bank Investment Credit Market

Supply
$$LRCINVRP_t^S = c_0 + c_1 LY_t + c_2(RINV_t - R1_t) + c_3 SD9903 \quad (3)$$

Demand
$$LRCINVRP_t^D = d_1 LY_t + d_2 RINV_t \quad (4)$$

where it is expected that $c_1>0$, $c_2>0$, $c_3<0$, $d_1>0$, and $d_2<0$.

Unit Root Test with Two Structural Breaks

This subsection investigates the degree of integration of the variables of interest. It is widely known that macroeconomic series often experience various breaks in their realizations. This is especially true for transition and emerging market economies, which often suffer from shocks due to radical policy changes or crises. Vogelsang and Perron (1998), through simulations, find that the unit root test size is sensitive to structural breaks and hence ignoring these breaks in the model specification may weaken the test power, thereby resulting in a misleading conclusion about the unit root hypothesis. Therefore, a unit root test whose size is invariant to the presence of structural breaks is needed[4].

This study implements a unit root test that allows for two structural breaks whose dates are determined endogenously. The procedure to be adopted is the one proposed by Lumsdaine and Pappell (1997) (henceforth LP), which is an extension of Zivot and Andrews (1992) or Perron (1997). Using innovational outlier (IO) framework[5], LP procedure is a modified version of the ADF test augmented by two endogenous breaks in both the time trend and the intercept. The model is written as follows:

$$\Delta y_t = \hat{\mu} + \hat{\alpha} y_{t-1} + \hat{\beta} t + \hat{\theta} DU1_t + \hat{\gamma} DT1_t + \hat{\omega} DU2_t + \hat{\psi} DT2_t + \sum_{i=1}^{k} \hat{c}_i \Delta y_{t-i} + \hat{e}_t . \quad (5)$$

where $DU1_t$ = 1 if $t>TB1$ and zero otherwise, $DU2t$ = 1 if $t>TB2$ and zero otherwise, $DT1t = t\text{-}TB1$ if $t>TB1$ and otherwise zero, and $DT2t = t\text{-}TB2$ if $t> TB2$ and otherwise zero.

into a full-fledged financial crisis until the mid 1998 and started to disrupt the credit market in early 1999.

[4] For excellent surveys of unit root studies see Phillips and Xiao (1998) and Maddala and Kim (2003).

[5] Perron (1989) considers the 1929 Great Crash as an example of structural break that occurred gradually because it lasted several years and hence assuming the DGP is of innovation outlier (*IO*), while the 1973 oil price shock as a break that occurred instantly. Accordingly, he modelled these two cases differently by applying *IO* to the former and *AO* to the latter in accordance with the DGP.

While the two dummy variables *DU1t* and *DU2t* capture structural changes in the intercept at time *TB1* and *TB2*, respectively, the other two dummy variables *DT1t* and *DT2t* capture shifts in the trend variable at time T_b1 and T_b2, respectively. The optimal lag order (k) is determined based on the *t*-Sig method in which the maximum lag order ($k$max) is set at 12. Further, the break dates (*TB1* and *TB2*) are estimated by minimising the value of the *t* statistic for α.

Bounds Test

Since the results of both unit root tests shown below suggest that the variables are of mixed degree of integration, the use of Johansen's procedure of identifying cointegration among variables is not appropriate. To mitigate this, another procedure, called bounds testing of cointegration based on an autoregressive distributed lag (ARDL) framework proposed by Pesaran *et al*. (2001), is attempted. The main advantage of the bounds testing procedure is that it allows us to implement cointegration test regardless of whether the underlying variables are *I(0)*, *I(1)*, or fractionally integrated. Thus, unit root tests are actually not required, except when inconclusive results (to be explained later) emerge.

The error correction representation of the ARDL model for equations 1, 2, 3 and 4 are as follows:

Working Capital Credit Market
Supply:

$$\begin{aligned}\Delta LRCWCRP_t = \; & a_0 + \sum_{i=1}^{k1} b_i \Delta LRCWCRP_{t-i} \\ & + \sum_{i=1}^{k2} c_i \Delta LY_{t-i} \\ & + \sum_{i=1}^{k3} d_i \Delta (RWC_{t-i} - R1_{t-i}) \\ & + \delta_1 LRCWCRP_{t-1} + \delta_2 LY_{t-1} \\ & + \delta_3 (RWC_{t-1} - R1_{t-1}) \\ & + \delta_4 SD9903_{t-1}. \qquad (6)\end{aligned}$$

Demand:

$$\begin{aligned}\Delta LRCWCRP_t = \; & \sum_{i=1}^{k1} b_i \Delta LRCWCRP_{t-i} \\ & + \sum_{i=1}^{k2} c_i \Delta LY_{t-i} + \sum_{i=1}^{k3} d_i \Delta RWC_{t-i} \\ & + \gamma_1 LRCWCRP_{t-1} \\ & + \gamma_2 LY_{t-1} + \gamma_3 RWC_{t-1} \\ & + \gamma_4 SD9903_{t-1}. \qquad (7)\end{aligned}$$

Investment Credit Market
Supply:7

$$\begin{aligned}\Delta LRCINVRP_t = a_0 & + \sum_{i=1}^{k1} b_i \Delta LRCINVRP_{t-i} \\ & + \sum_{i=1}^{k2} c_i \Delta LY_{t-i} \\ & + \sum_{i=1}^{k3} d_i \Delta (RINV_{t-i} - R1_{t-i}) \\ & + \lambda_1 LRCINVRP_{t-1} + \lambda_2 LY_{t-1} \\ & + \lambda_3 (RINV_{t-1} - R1_{t-1}) \\ & + \lambda_4 SD9903_{t-1}. \qquad (8)\end{aligned}$$

Demand

$$\begin{aligned}\Delta LRCINVRP_t = & \sum_{i=1}^{k1} b_i \Delta LRCINVRP_{t-i} \\ & + \sum_{i=1}^{k2} c_i \Delta LY_{t-i} + \sum_{i=1}^{k3} d_i \Delta RINV_{t-i} \\ & + \beta_1 LRCINVRP_{t-1} + \beta_2 LY_{t-1} \\ & + \beta_3 RINV_{t-1} + \beta_4 SD9903_{t-1} \qquad (9)\end{aligned}$$

where Δ is difference operator, the order of ARDL is (*k1*, *k2*, *k3*) so that $i = 1...k1$; $i = 1...k2$; and $i = 1...k3$.

The null hypotheses of no cointegration, defined by $H_0: \delta_1 = \delta_2 = \delta_3 = \delta_4 = 0$; $\gamma_1 = \gamma_2 = \gamma_3 = \gamma_4 = 0$; $\lambda_1 = \lambda_2 = \lambda_3 = \lambda_4 = 0$; and $\beta_1 = \beta_2 = \beta_3 = \beta_4 = 0$, are tested against their alternatives, which are $H_1: \delta_1 \neq 0,\ \delta_2 \neq 0,\ \delta_3 \neq 0,\ \delta_4 \neq 0$; $\gamma_1 \neq 0,\ \gamma_2 \neq 0,\ \gamma_3 \neq 0,$ $\gamma_4 \neq 0$; $\lambda_1 \neq 0,\ \lambda_2 \neq 0,\ \lambda_3 \neq 0,\ \lambda_4 \neq 0$; and $\beta_1 \neq 0,\ \beta_2 \neq 0,\ \beta_3 \neq 0,\ \beta_4 \neq 0$, by means of the *F*-test. Since the asymptotic distributions of the *F*-statistic are non-standard irrespective of whether the variables are *I(0)* or *I(1)*, Pesaran *et al*. (2001) calculate two sets of asymptotic critical values. While the first set assumes that all variables are *I(0)* and serves as lower bounds, the second assumes they are all *I(1)* and serves as upper bounds. If the computed *F*-statistic falls below the lower bound the null hypotheses cannot be rejected. Conversely, if it falls above the upper bound, then the null hypotheses can be rejected. However, if it falls between these two bounds, the result is inconclusive and a unit root test is required to determine the order of integration of the variables. Once cointegration is confirmed, the next stage is estimating the long-run coefficients of both bank lending supply and demand relations and their associated ARDL-based error correction models.

Data Description

In this study monthly data are employed starting from the first month of 1985 to the last month of 2007, covering a total of 276 observations. All the data except real GDP (*Y*) are available in a monthly frequency. Therefore, the frequency of *Y* series has been converted from quarterly into monthly by using the distributive method. The precise data sources are given in Table 1 and the series are plotted in Figure 1.

RESULTS DISCUSSION

Unit Root Test with Two Structural Breaks

For the sake of comparison the results of the traditional ADF unit root test are reported in Table 2. Based on the AIC model selection criterion for determining the optimal lag lengths, the unit root null hypothesis is rejected at 5 per cent level for 3 series (*R1*, *RWC* and *RINV*). Table 3 reports the results of the LP test procedure. The critical values for α are taken from LP (1997) for model CC. As expected, this procedure can reject the null hypothesis in favour of stationary process at 5 per cent level for far higher number of series (5) than the traditional procedure. Accordingly, these 5 series are considered as integrated of degree zero when two structural breaks are allowed. Taken together, based on the unit root tests the 6 variables in the Indonesian bank credit markets are of mixed degree of integration.

Table 1: Description and Sources of Data

No	Variable	Description	Sources
2	R1	Interbank call-money rate	IFS-IMF
4	RWC	Bank working capital landing rate	IFS-IMF
5	RINV	Bank investment landing rate	IFS-BI
6	RCWCRP	Real bank working capital rupiah loan	IFS-BI
7	RCINVRP	Real bank investment rupiah loan	IFS-BI
12	CPI	Consumer Price Index (2000=100)	IFS-IMF
13	Y	Real Gross Domestic Product (2000 =100)	BPS

BPS = Badan Pusat Statistik (Central Bureau of Statistics) Indonesia

IFS-IMF = International Financial Statistics – International Monetary Fund (IMF) (CD-ROM database)

IFS-BI = Indonesian Financial Statistics – Bank Indonesia (Published monthly and at < http://www.bi.go.id >)

MSCI = Morgan Stanley Capital International Inc, available at <http://www.msci.com/equity/index2.html>.

Table 2: ADF Unit Root Test

No	Series	TI/I[1]	$t_{\hat{\alpha}}$	S/N
1	R1	I	-3.5975*	S
2	RWC	I	-3.1683*	S
3	RINV	I	-3.1372*	S
4	LRCWCRP	I	-2.1481	N
5	LRCINVRP	I	-1.7599	N
6	Y	I	-1.9647	N

Notes:
1. [1] *TI* and *I* denote the regression where the trend function includes both time-trend and intercept and intercept only, respectively (the inclusion of *T* or *I* is determined by the result that most likely rejects the unit root null hypothesis).
2. * and ** denote the rejection of the unit root hypothesis at 5% and 10% level, respectively.
3. S and N denote stationary and nonstationary respectively.

Table 3: Unit Root Test with Two Breaks at Unknown Dates

No	Series	*TB*1	*TB*2	*t*-statistic for α	Lag	S/N
1	R1	1993M06	1997M08	-7.97851*	12	S
2	RWC	1990M08	1997M08	-9.04684*	12	S
3	RINV	1993M08	2000M01	-5.87939	6	N
4	RCWCRP	1997M05	1999M03	-8.63096*	5	S
5	RCINVRP	1997M02	1999M03	-7.83381*	0	S
6	Y	1997M05	1998M07	-8.14886*	12	S

Notes:
1. The critical values at 1%, 2.5%, 5%, and 10% levels of significance are 7.34, 7.02, 5.82, and 5.49, respectively (Lumsdaine and Papell, 1997);
2. * and ** denote rejection of the unit root hypothesis at 5% and 10% levels, respectively.

Table 3 also reports the estimated break dates for each series. For the estimated *TB*1, the dates coincide with the surge of foreign capital inflow in 1992, which continued until the mid 1997 before the 1997 currency crisis started. As for the estimated *TB*2, the dates mostly point to the financial crisis that started to unfold in August 1997 and did not recede until the end of 2000. Of particular importance is the estimated break date for *LRCWCRP* and *LRCINVRP* which are used to construct the shift dummy $SD99DU_t$ that has zero for observations before the third month of 1999 and one thereafter. Figure 1 depicts the plot for each series along with its respective estimated break dates.

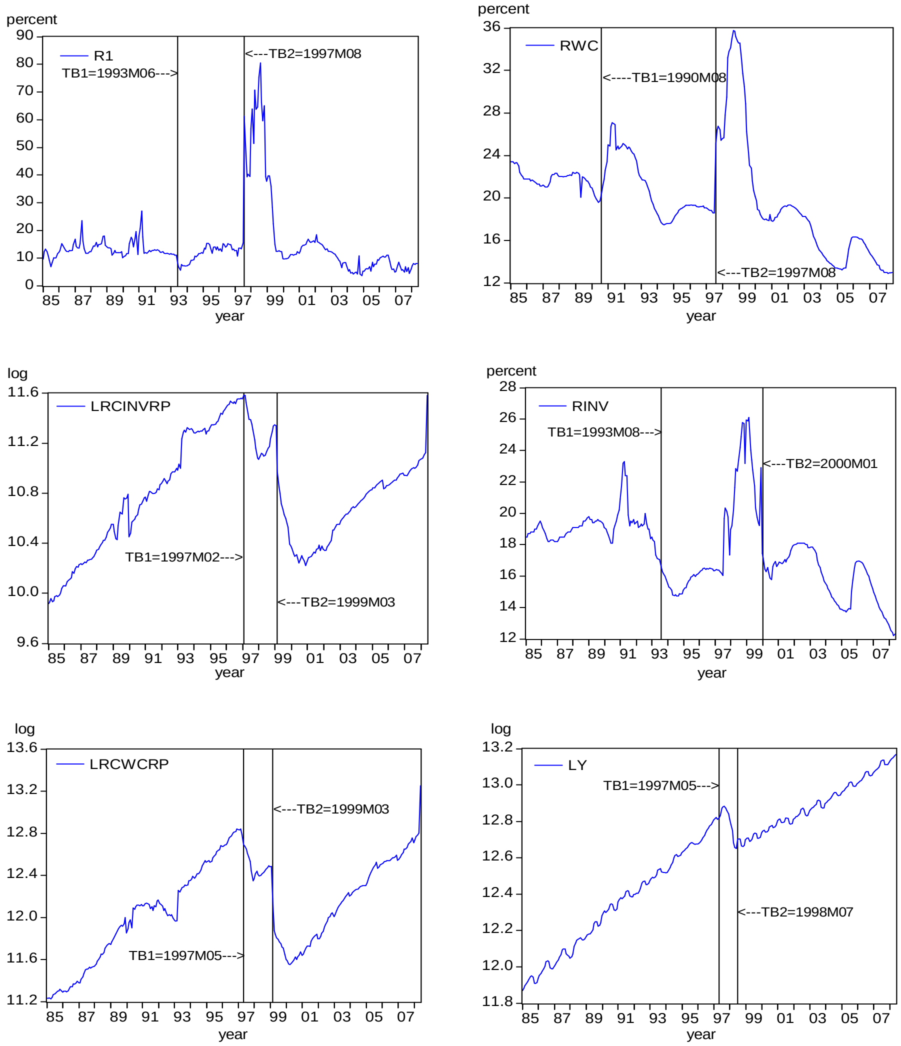

Figure 1: Series with Two Endogenously-Determined Break Dates (TB1 and TB2)

Bounds Test

Before testing the null hypotheses of no cointegration for equations (6) to (9), the order of lags on the first-differenced variables must be determined. Since the test results usually vary with the order of lags, different orders of lags: 2, 4, 6, 8, 10, and 12 on the first difference of each variable, are attempted. The computed *F*-statistic for each order of lag in each equation associated with the working capital credit and investment credit markets is reported in Tables 4 and 5 respectively. The test outcome turns out not to vary significantly with the choice of lag order. The only exceptions are the results for the supply of working capital credit (*LRCWCRP*) relation specified with the lag orders 2 and 4, namely ARDL (2, 2,

2) and ARDL (4, 4, 4) respectively, and for the demand for investment credit (*LRCINVRP*) specified with the lag order 12, namely ARDL (12, 12, 12). The computed *F*-statistics resulting from these three exceptions are inconclusive. As reported in the tables, except for these three, all orders of lags for both relations of respective markets provide *F*-statistic, which is significant at 5 per cent. Therefore, the results largely suggest the existence of long-run supply and demand equations in both credit markets.

Table 4: Bounds F-Statistics (Working Capital Credit Market)

Supply of Working Capital Credit		Demand for Working Capital Credit	
Order of Lag	F-Statistics[a]	Order of Lag	F-Statistics[b]
2	3.206	2	12.616**
4	3.779*	4	10.477**
6	4.284**	6	12.774**
8	6.452**	8	12.055**
10	4.934**	10	7.281**
12	4.088**	12	4.984**

Notes:

1. [a] The critical value bounds are given in Table CI.iii (with an unrestricted intercept and no trend; number of regressors = 4), Pesaran et al. (2001). They are 2.86 – 4.01 at the 5 per cent significance level and 2.45 – 3.52 at the 10per cent significance level.
2. [b] The critical value bounds are given in Table CI.i (with no intercept and no trend; number of regressors = 4), which are 2.26 – 3.48 at 5per cent significance level and 1.90 – 3.01 at 10per cent significance level.
3. * denotes that the F-statistic falls above the 10 per cent upper bound and ** denotes above the 5 per cent upper bound.

Table 5: Bounds F-Statistics (Investment Credit Market)

Supply of Investment Credit		Demand for Investment Credit	
Order of Lag	F-Statistics[a]	Order of Lag	F-Statistics[b]
2	5.079**	2	5.656**
4	5.402**	4	6.181**
6	5.637**	6	5.791**
8	5.271**	8	5.933**
10	7.666**	10	4.975**
12	5.025**	12	2.444

Notes:

1. [a] The relevant critical value bounds are given in Table C1.iii (with an unrestricted intercept and no trend; number of regressors = 4), Pesaran *et al*. (2001). They are 2.86 – 4.01 at the 95 percent significance level and 2.45 – 3.52 at the 90per cent significance level.
2. [b] The critical value bounds are given in Table CI.i (with no intercept and no trend; number of regressors = 4), which are 2.26 – 3.48 at 5 per cent significance level and 1.90 – 3.01 at 10 per cent significance level.
3. * denotes that the F-statistic falls above the 10 per cent upper bound and ** denotes above the 5 per cent upper bound

Bank Lending Identification

The next stage is estimate equations 1 to 4 using the lag selection criteria AIC, SBC, and the maximum lag order of 12[7]. Thus, each equation is estimated three times. Tables 6 and 7 report the estimation results for both credit markets whose lag orders are determined based on AIC. Note that the estimated coefficients obtained from all three model selection criteria are quite similar and all regressors, except the interest rate spread in the working capital credit market *(RWC-R1)* estimated using maximum lag order (12), are highly significant. As expected, all estimated coefficients are correctly signed.

Table 6: Estimated Cointegration Relations based on ARDL (Working Capital Credit Market)

Supply ARDL (12, 10, 0) Based on AIC	$L\ RCWCRP_t^S = -6.421 + 1.756\ LY_t + 0.007\ (RWC_t - R1_t) - 0.875\ SD9903$ (-7.479) (20.730) (2.360) (-16.318)	(10)
Short-run adjustment coefficient, **α**	= **-0.193** (-6.501)	
Demand ARDL (6, 0, 2) Based on AIC	$L\ RCWCRP_t^D = 1.305\ LY_t - 0.079\ RWC_t - 1.004\ SD9903_t$ (29.980) (-4.239) (-7.426)	(11)
Short-run adjustment coefficient, **α**	= **-0.096** (-5.270)	

Note: Figures in parentheses are *t* statistic.

Table 7: Estimated Cointegration Relations based on ARDL (Investment Credit Market)

Supply ARDL (1, 3, 0) Based on AIC	$L\ RCINVRP_t^S = -7.076 + 1.802\ LY_t + 0.006\ (RINV_t - R1_t) - 1.148\ SD9903$ (-10.751) (24.275) (3.584) (-24.152)	(12)
Short-run adjustment coefficient, **α**	= **-0.194** (-7.852)	
Demand ARDL (1, 0, 0) Based on AIC	$L\ RCINVRP_t^D = 1.140\ LY_t - 0.072\ RINV_t - 1.043\ \mathrm{SD9903}_t$ (35.498) (-4.213) (-11.377)	(13)
Short-run adjustment coefficient, **α**	= **-0.092** (-6.026)	

Note: Figures in parentheses are *t* statistic.

All estimated coefficients in the relations are correctly signed and statistically significant. The estimated income elasticity of demand for working capital loans *(RCWCRP)* and income elasticity of demand for investment loans *(RCINVRP)* are 1.31 and 1.14, respectively, which are smaller than that found in Agung et al. (2002) and Kakes (2000)[6]. Although higher than those found in Agung et al. (2002), the estimated interest rate elasticities, -1.71 and -1.24 respectively, offer similar intuition[7]. That is, the demand for the working capital loans is more sensitive to the loan rate than the demand for investment loans since the latter is long-term loans which are more likely subject to long-term contract. As expected the estimated coefficients on the shift dummy *SD9903* are correctly signed and statistically significant in both supply and demand equations in both credit markets. Because the shift dummy is present in both supply and demand relations and is statistically significant, the financial crisis have negative impact on both bank lending supply and demand. The decline in demand results from the fact that during the financial crisis many companies suddenly technically went bankrupt and had to cancel many new investment plans and terminate numerous investment projects in progress. Nonetheless, comparing the estimated coefficient on the shift dummy in these two relations reveals that it is greater in magnitude and statistically more significant in the supply relation than in the demand relation. At least it can be said that the negative impact of the financial crisis on the supply of investment credit is more pronounced than on the demand side. The negative sign of its coefficient may be interpreted as evidence for the existence of credit crunch following the financial crisis[8]. That is the financial crisis has significantly reduced the quantity of bank loans supply.

The error correction representation of ARDL models provide the estimated short-run adjustment coefficients which suggest that in the short run the market for working capital and investment credits is dominated by supply rather than demand. As also reported in Tables 6 and 7, the corresponding ~~short-~~ short term adjustment coefficients indicate that bank loans adjust significantly in the direction of both long-run supply and demand equations. However, comparing the magnitude and statistical significance, the adjustment to the supply equation is greater and statistically more significant than the adjustment to the demand equation. This is true of both markets. The estimated coefficients are -0.193 and -0.096 respectively in the working capital loans market and -0.194 and -0.092 respectively in the investment loans market. This suggests that in the short run the market for working capital and investment credits is dominated by supply rather than demand.

CONCLUSIONS

This paper has attempted to identify the existence of bank landing channel of monetary policy transmission by developing an error correction representation of the autoregressive distributive lag (ARDL) model of the Indonesian bank credit market. Two separate models have been estimated: the working

[6] Agung et al. (2002) found the income elasticity of demand for working capital loans 1.8 and for investment loans 2.7. Kakes (2002) found for the elasticity for credit market as a whole 1.76.

[7] The interest rate elasticity of demand for bank credit is the product of estimated coefficient on the lending rate (semi elasticity of the rate) and the sample mean of that rate. Thus, the working capital rate elasticity is (-0.079×21.65) = -1.71 and the investment credit rate is (-0.072×17.41) = -1.24. Agung *et al.* (2002) found the interest rate elasticities are -0.36 and -0.20, respectively.

[8] An attempt is made not to impose a restriction on the shift dummy in the demand relation for both market models, but the estimated coefficients are not significant. Instead, when an exclusion restriction is imposed the result improves statistically.

capital credit market and investment credit market models. Departing from previous studies, each model takes account of one structural break associated with the 1998 financial crisis. The exact date of the crisis has been endogenously determined by implementing a unit root test that allows for two structural breaks. Some of the estimated break dates coincide with the 1998 financial crisis. Since the variables of the Indonesian bank credit markets are of mixed degree of integration, as suggested by the unit root test, bounds test procedure has been implemented, instead of Johansen's cointegrating procedure. Variables in both bank lending supply and demand equations turn out to be cointegrated.

The estimated equation for bank loans in the error correction model produces a result that bank loans adjust more strongly in the direction of the supply equation. This is true for both markets. This suggests that in the short run both loans markets are dominated by supply rather than demand. Likewise, a monetary-policy-induced disturbance in the equilibrium quantity of bank loans more likely originates from the supply side than the demand side. Similarly, the results also indicate the existence of following the financial crisis.

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