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Credit Markets and the Propagation of Korea's 1997 Financial Crisis

Peter L. Rousseau* and Jong Hun Kim†

We consider the roles of monetary shocks and tightening credit market conditions in the transmission of South Korea's 1997 financial crisis to the real sector, and compare the relative impacts of these factors on production in light and heavy industries. Using structural regression equations, vector autoregressive models, and the accompanying dynamic forecasts, we find that the ratio of commercial bills dishonored to the total value of bills to be cleared can explain the decline in industrial production more fully than either the decline in the real stock of money or the spread between yields on corporate and government bonds. These results are most emphatic in light industry, for which small and medium-sized firms account for more than 70% of the total value added. Since fluctuations in the dishonored bills ratio may reflect components related to increases in the cost of credit intermediation and its effect on small and medium-sized firms more precisely than the corporate–government bond spread, we interpret the evidence as suggestive of a credit channel and "flight to quality" at work.

JEL Classification: E5, F4

1. Introduction

In 1997, the East Asian economies faced a sudden financial crisis. Three countries, Indonesia, Korea, and Thailand, received emergency loans from the IMF. The rapid decline in economic activity that occurred in the wake of speculative attacks at that time has led to no shortage of explanations for the crisis among policy makers. According to the IMF, for example, fundamental weaknesses in the financial intermediaries of the crisis economies were largely at fault. Radelet and Sachs (1998), on the other hand, contend that the crisis was a self-fulfilling one with roots in the inherent instability of international financial markets. In either case, what does seem clear is that the crisis involved changes in credit market conditions that reflected more than purely monetary phenomena. In this paper, we seek to further our understanding of the crisis in South Korea by quantifying the importance of such nonmonetary factors in precipitating a credit "crunch" that acted most emphatically through small and medium-sized enterprises. 1

Our methodology includes structural regression models that use versions of the Lucas (1972) supply equation modified to include indicators of credit market conditions. We then extend the analysis to consider the longer-term effects of money, credit conditions, and real

* Department of Economics, Vanderbilt University, Box 1819 Sta. B, Nashville, TN 37235, USA; National Bureau of Economic Research; E-mail peter.l.rousseau@vanderbilt.edu; corresponding author.

† Department of Economics, Vanderbilt University, Box 1819 Sta. B, Nashville, TN 37235, USA; E-mail jong.h.kim@vanderbilt.edu.

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exchange rates on industrial production with a series of vector autoregressive (VAR) systems. The yield spread between corporate and government bonds, and the ratio of dishonored commercial bills to the total value of bills to be cleared, serve as indicators of the state of the credit market. We then assess the relative abilities of these variables to explain the length and depth of Korea's financial crisis across heavy and light industry.

We find that increases in yield spreads and the dishonored bills ratio, whether driven by increases in business risk or lowered expectations among investors about the future of the Korean economy, had effects that extended well beyond shifts in simple precautionary and speculative demands for the won. We also find that these effects were strongest for light industry, in which small and medium-sized firms accounted for more than 70% of value added in 1999, and that the dishonored bills ratio explains more of the economic decline than the yield spread in our econometric models. We interpret these findings as consistent with the operation of a mechanism much like a "credit channel" and associated "flight to quality" through which increases in risk and the cost of credit intermediation cause the brunt of a credit crunch to fall disproportionately on smaller firms. Though resembling a credit channel in its effects, however, our story differs from the standard treatment because events did not begin with routine policy intervention by the Bank of Korea, which operated exclusively through bank credit.

2. Background

At the end of 1997, Korea experienced its first financial crisis, which resulted in the nation's worst economic performance in 40 years. Real GDP growth was −6.7% in 1998, which stands in sharp contrast to the 7.7% growth achieved during the "miracle" years of 1960 to 1996.[1] Moreover, in 1998 the unemployment rate reached 6.8%, up from only 2.6% the year before. The previous postwar high of 7.1% had been reached more than 30 years earlier in 1964. It is clear that the financial crisis coincided with a period of severe economic stress.

Whether the crisis was anticipated or not is crucial to investigating its causes. Krugman (1998) argues that it was indeed anticipated, and that foreign investors expected to be bailed out. If this were the case, however, foreign creditors would have disinvested upon observing negative signals about the future of the economy, and this does not appear to have been the case. Further, there is little empirical evidence that the crisis was anticipated. Table 1 summarizes economic conditions in 1997. Foreign direct investment (FDI) to Korea was on the rise. Over the first nine months of 1997, there was 50% more FDI than there had been over the corresponding period in 1996. And even though the current account balance showed a large deficit due to worsening terms of trade, net capital inflows had up to then been able to cover most of the deficit. Overall, there was a shortfall of only US$0.5 billion until October. The dishonored bills ratio was relatively low until the crisis, but then rose nearly sevenfold over the next three months.

The crisis seemed to stem from the failure of small investment banks. When the capital market opened in the 1990s, the Korean government granted new business permits to more than 20 such banks—permits that included the right to borrow abroad. By 1997, some of these

[1] The lowest previous annual growth rate since 1960 was −2.1%, which occurred in 1980 in the midst of the second oil shock and a climate of political instability.

Table 1. Economic Conditions for South Korea in 1997

	Q1	Q2	Q3	Oct.	Nov.	Dec.
Changes in terms of trade (%)	−8.39	2.51	−0.00	0.08	−6.17	−4.92
Current account balance[a]	−7.35	−2.72	−2.05	−0.49	0.86	3.59
Capital account balance[a]	4.04	6.57	0.62	0.92	−4.46	−6.37
Foreign direct investment[a]	0.62	0.79	0.61	0.15	0.29	0.39
Dishonored bills ratio (%)	0.29	0.30	0.32	0.56	0.48	2.09

[a] In billions of U.S. dollars.

new banks had already experienced liquidity problems due to maturity mismatches and bad investments, and in October the ratio of loans in default exceeded 5% for five of these banks. As they tried to obtain more U.S. dollars to repay short-term debts, changing expectations about the path of exchange rates increased the demand for U.S. dollars at the worst possible time. Kaminsky and Reinhart (1999) describe the type of downward spiral that ensued. The government intervened to protect the won, yet the currency continued to weaken as foreign reserves at the Bank of Korea came under increasing pressure. Finally, with its reserves drained, the Korean government requested an emergency loan from the IMF on October 21.

Even after the emergency loan was approved on December 3, however, the won continued to fall and did not stabilize until January 28, 1998, when foreign creditors agreed to roll over the short-term debt of many Korean financial intermediaries. Interestingly, the solution came from market interactions between debtors and creditors rather than intervention by the Korean government or the emergency loan. The difficulties involved with arriving at a monetary solution motivate us to consider nonmonetary factors, such as those driven by a severe credit crunch, in explaining the length and depth of the economic downturn.

3. Credit Market Conditions and Real Activity

Monetary theory often focuses on two main channels through which a shift in the supply of money can affect real economic activity. The first is the traditional money view, in which changes in the money supply affect the demand for household consumption and business investment through the interest rate. The second is the lending or "credit" channel, in which a decrease in the money supply causes a reduction in the supply of loans to the private sector. For the latter to operate, it is necessary that: (i) banks adjust their loan supply in response to monetary shocks, (ii) loans are not perfect substitutes on a bank's balance sheet, and (iii) there is imperfect substitutability for firms between bank loans or other commercial bills and bond issues (Bernanke and Blinder 1988). When these conditions are met, interest rates on loans increase relative to those on other securities, and the widening of the spread between risky and safer financial assets reflects the extent to which monetary policy has become more restrictive. An active credit channel can also generate a flight to quality in which banks choose to lend to larger and safer borrowers at times of monetary stringency (Bernanke, Gertler, and Gilchrist, 1996). This means that some firms, particularly smaller ones, end up relying more on existing collateral and internal funds to support their operations. Our contribution extends the idea of a lending channel to analyze a financial crisis in which the real sector is not so much affected by a declining supply of money as by increases in the costs of providing financial intermediation services, and how this might generate a flight to quality in the open market for credit.

Recent studies have taken varied approaches to examining the credit crunch in Korea, with most focusing on the bank-lending channel. Using commercial banks' data, for example, Ferri and Kang (1998) found that less well-capitalized banks tended to increase their lending rate and reduce loans during the crisis. Consistent with this, Kim (1998) found a large excess demand for bank loans after the crisis. In contrast, Ghosh and Ghosh (1999) found little evidence of an excess demand for credit. Borensztein and Lee (2002) analyze the credit crunch using firm-level data and conclude that firms affiliated with Korea's largest informal business networks, the "chaebols," appear to have lost some of the preferential access to credit that they enjoyed in the precrisis period, suggesting that the credit crunch affected a wide range of firms. Ding, Domac, and Ferri (1998) link the credit crunch to sharp increases in yield spreads between risky and risk-free assets. Investigating five different countries, Indonesia, Korea, Malaysia, Philippines, and Thailand, they observe that these effects were significant only for Korea and Malaysia.

We suspect that the mixed findings of earlier studies may to some extent reflect limitations of restricting analysis to a more traditional credit channel framework. Even though the credit view has proven to be an effective tool for explaining the effects of monetary policy on real activity, it is indeed focused on policy rather than the systematic failure of a financial system. In the case of Korea in 1997, we believe that an analysis of nonmonetary factors, defined more generally than the credit channel to include supply-side factors both inside and outside of the banking system, may shed additional light on our understanding of the length and depth of the crisis. In an important earlier study along these lines, Bernanke (1983) makes a case for such nonmonetary factors in the propagation of the Great Depression in the United States.[2]

4. Measures of Credit Market Conditions

When banks make loans, they incur screening, monitoring, and accounting costs, as well as losses from defaulting borrowers. Since lending is an enterprise that is generally believed to involve scale economies, it is likely that, on an amount-loaned basis, the expected costs are higher for smaller, less established firms than for larger ones. One might then expect that an increase in the cost of credit intermediation would reduce lending overall, and particularly to smaller firms. Ideal summary measures of the cost of lending, however, are difficult to identify for empirical testing. One candidate, for example, could be the average interest rate on bank loans, but even this would not be entirely satisfactory because it reflects loans actually made rather than the shadow cost of borrowing. Indeed, if banks tend to make only the safest loans in a period of retrenchment, loan rates may even move inversely to the cost of credit intermediation.

Given these challenges, our analysis starts by using the interest rate between a safe and riskier bond as a measure of general credit conditions. Unfortunately, only the spread between

[2] Bernanke found that adding indicators of a general financial crisis such as the deposits of failed banks and the liabilities of failed businesses improved upon a purely monetary explanation of output fluctuations, reducing the mean square error in simulations of structural output equations by about 50%.

corporate and government bonds is available to us for Korea over the full period of our study.[3] We use it even though corporate bond yields reflect interest rates that are available to relatively large and creditworthy firms, thus making the spread understate the true difference in yields between representative risky and safe assets. If these differences are greatest in times of financial crisis and credit contraction, and banks prefer to lend to larger and safer firms at such times, then the corporate–government spread is likely to explain less of the real effects of a crisis than a measure that more closely reflects lending frictions associated with smaller businesses.[4] Despite this limitation, previous studies have found that similar "spread-type" variables predict real output quite well. Friedman and Kuttner (1993), for example, show a link between the spread on six-month prime-rated commercial paper and 180-day Treasury bills and subsequent fluctuations in real economic activity in the United States over the postwar period.

We expect the spread to widen as a financial crisis deepens and credit risks become increasingly difficult to evaluate, yet a portion of the spread's explanatory power is likely due to pure financial market anticipations of future declines in real activity. A variable that more closely reflects business defaults and failures, such as the ratio of dishonored commercial bills to the total value of bills to be cleared, likely reflects past business conditions more than anticipations of future declines, and thus would appear to be a good instrument for purging the spread of its anticipatory component.[5] If the instrumented spread can still predict output well, this would be suggestive of a credit channel at work.

Increases in the dishonored bills ratio itself may also signal a riskier climate for conducting business and a rise in the cost of credit intermediation, and we would in this case expect increases in the ratio to affect real activity through the provision of credit. Indeed, commercial bills often wind up in default because the issuing firm is unable to obtain the short-term funding needed to meet obligations as they become due. These frictions are heightened in a financial crisis as banks seek borrowers with low costs of screening and monitoring. For this reason, we also use the dishonored bills ratio directly in our output regressions as a measure of credit conditions. Since many of Korea's smaller firms have direct access to standard commercial bills that provide much of the working capital to the business sector, the dishonored bills ratio might even be preferable to the corporate–government yield spread in capturing the credit conditions faced by small and medium-sized firms.

Of course, business failures, as reflected in the dishonored bills ratio, could cause temporary resource misallocations as businesses take time to wind up their affairs. To the extent that such resources (i.e., physical capital, labor, and management) remain temporarily idle or production inflexibilities prevent their reallocation to firms that could deploy them more effectively, output may fall. This means that some of the explanatory power of the dishonored bills ratio could reflect real-side shocks rather than financial factors. Indeed, if this channel were to dominate, a strong role of the ratio in predicting output would be only suggestive of a credit crunch. To reduce the influence of these possible real-side effects on estimated coefficients for the nonmonetary financial measures, our empirical models always include past values of industrial production as explanatory variables.

[3] We compute the spread use monthly yields on 3-year corporate bonds and 3-year Treasury bonds from the over-the-counter market as reported by the Bank of Korea.

[4] For example, if it were available, the yield differentials between typical BBB- and AAA-rated bonds or between CCC- and BBB-rated bonds might reflect the cost of credit intermediation more closely than the spread that we use.

[5] We construct the dishonored bills ratio using monthly data from the Bank of Korea.

To quantify and compare the effects of the financial crisis on light and heavy industries, we consider the impact of our nonmonetary financial indicators on these sectors separately as well as on total industrial production. Since light industry is dominated by small firms, if a flight to quality was at work, we would expect to find production in light industry affected more sharply than in either heavy industry or the aggregate.

5. Structural Analysis

Our econometric analysis uses monthly data from January 1987 to December 2003. We begin in 1987 since this was the first year for which the Bank of Korea reported yields on corporate and government bonds. It also covers the period of the financial crisis and an adequate period before and after. The monthly frequency allows us to focus on the short-run effects of the crisis. Specifically, we use the industrial production index (1995 = 100) from Korea's National Statistical Office,[6] and the producer price index (1995 = 100) and money stock (*M2*) from the Bank of Korea.[7] All series are seasonally adjusted. We first examine the effects of changing credit market conditions on aggregate industrial production, and then re-estimate for heavy and light industry separately. We then generate dynamic forecasts of industrial production to compare the relative predictive abilities of our two credit variables. Finally, we compare both against a baseline that includes only monetary shocks.

Methodology

Following Lucas (1972), our structural models assume that only unanticipated changes in money or prices affect real output, and so we begin by extracting the unexpected components of changes in the money supply from the *M2* series. To do this, we use the two-step method described in Barro (1978). The first step identifies the monetary shocks as residuals from the equation

$$\dot{M}_t = a + \sum_{i=1}^{4} b_i \dot{Y}_{t-i} + \sum_{i=1}^{4} c_i \dot{P}_{t-i} + \sum_{i=1}^{4} d_i \dot{M}_{t-i} + e_t, \tag{1}$$

where $\dot{Y}$ is the growth rate of the industrial production index, $\dot{P}$ is inflation in the producer price index, and $\dot{M}$ is the growth rate of the money stock (*M2*). We use a lag order of four as selected by the Schwartz criterion. Augmented Dickey–Fuller (ADF) and Phillips–Perron (1988) (PP) tests reject the null hypothesis of a unit root for all of these growth rates.

In the second stage, we use the residuals from the first stage to estimate the effects of credit conditions and the unanticipated monetary shocks on industrial production. The baseline specification, which considers only money, is

[6] This index represents the level of output produced in the secondary sector, which includes manufacturing (94.4%), electricity, gas and water (5.0%), and mining (0.6%).

[7] We use *M2* as our monetary variable instead of *M1* because *M2* has been the target for Korean monetary policy since 1979.

Table 2. Second-Stage Regression Results (Dependent Variable is Total Industrial Production *YT*)

	Baseline Model	Credit Variables (CCI)		
		SPR	CCI (IV)	DBR
Constant	0.458* (0.145)	0.493* (0.143)	0.499* (0.145)	0.468* (0.143)
$YT_{t\text{-}1}$	0.663* (0.069)	0.637* (0.068)	0.645* (0.069)	0.634* (0.068)
$YT_{t\text{-}2}$	0.220* (0.068)	0.238* (0.068)	0.230* (0.068)	0.246* (0.068)
$(M - M^e)_t$	0.361 (0.222)	0.419 (0.220)	0.413 (0.233)	0.575* (0.233)
$(M - M^e)_{t-1}$	−0.176 (0.223)	−0.101 (0.222)	0.002 (0.237)	−0.065 (0.224)
$(M - M^e)_{t-2}$	0.541* (0.222)	0.526* (0.219)	0.616* (0.223)	0.576* (0.219)
C_t		−0.443* (0.173)	−1.371* (0.662)	−2.314* (0.885)
Trend	0.001* (0.000)	0.001* (0.000)	0.001* (0.000)	0.001* (0.000)
F-test (*p*-value)				
Money shocks	0.727 (0.029)	0.844 (0.024)	1.031 (0.013)	1.087 (0.006)
Residual analysis				
ADF test	−7.730*	−14.597*	−14.636*	−14.707*
PP test	−14.550*	−14.733*	−14.773*	−14.840*

YT is the log of the industrial production index, *M* is unanticipated money shocks, SPR is the observed spread, DBR is the ratio of dishonored bills to the total amount of bills to be cleared, and CCI (IV) is the fitted value of the spread regressed on DBR. The optimal lags for the residual analysis were chosen using the Akaike information criterion. * Represents statistical significance at the 5% level. Standard errors are in parentheses.

$$Y_t = \alpha + \sum_{i=1}^{2} \beta_i Y_{t-i} + \sum_{i=0}^{2} \gamma_i (M - M^e)_{t-i} + \delta t + \varepsilon_t, \tag{2}$$

where Y is a log level of the industrial production index, $(M - M^e)$ is the unanticipated shock to money growth, and t is a linear time trend. This time the Schwartz criterion selects a model with two lags. Next, we add the credit variables (C) to the baseline as follows:

$$Y_t = \alpha + \sum_{i=1}^{2} \beta_i Y_{t-i} + \sum_{i=0}^{2} \gamma_i (M - M^e)_{t-i} + \xi C_t + \delta t + \varepsilon_t, \tag{3}$$

where C is the corporate–government bond yield spread (SPR), the spread instrumented with the dishonored bills ratio (CCI [IV]), or the dishonored bills ratio itself (DBR).[8,9]

Results from Structural Regressions for Aggregate Industrial Production

Table 2 shows the results from the second-stage regressions. The first column, which corresponds to the baseline model, indicates that unanticipated money shocks have a positive effect on aggregate industrial production. This is indicated in the row labeled "money shocks," where the algebraic sum of the contemporaneous and lag coefficients is positive and statistically significant at the 5% level. The results for specifications with the three measures of credit market conditions (C_t) entering one at a time appear in the next three columns. There we find

[8] In Table 4 of section 5 we report that ADF and PP tests reject the null hypothesis of a unit root for both the observed spread and the dishonored bills ratio.

[9] We also estimated models with multiple lags of the credit variables and obtained similar results. (i.e., lags of the credit variables were jointly significant in standard *F*-tests).

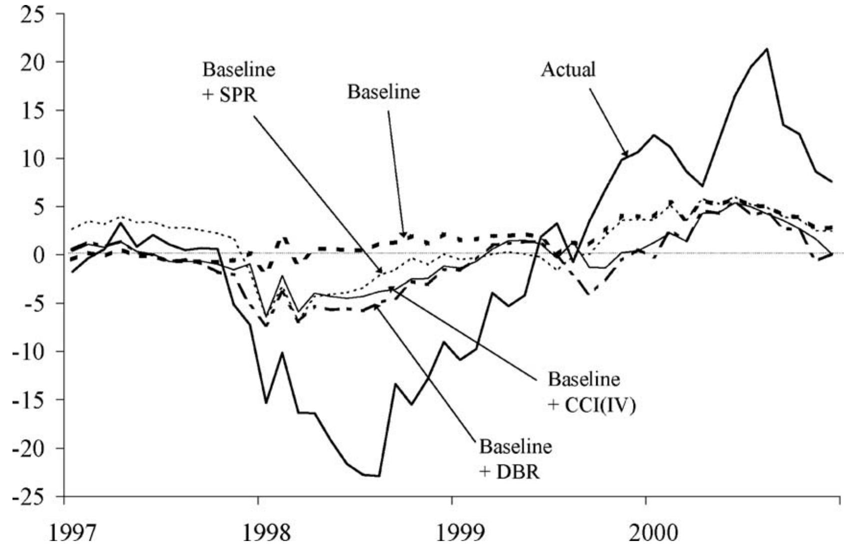

Figure 1. Actual Industrial Production and Dynamic Simulations

that the coefficients for all of the credit variables are negative and statistically significant at the 5% level. This suggests that industrial production did indeed respond to monetary shocks, but that a "nonmonetary" channel also played a key role in their propagation. A check of the residuals indicates that they are indeed stationary in all of the specifications.

To quantify the incremental contributions of unanticipated money shocks and the credit variables to real activity, we compute dynamic forecasts of industrial production on the basis of the coefficient estimates for the baseline model and each of the alternative specifications. Adding SPR, CCI (IV), and DBR reduces the mean-square forecasting errors by 36.5%, 42.7%, and 50.9%, respectively, compared with the baseline, from December 1997 to June 1999, which covers the period from the start of the crisis to the month when industrial production returned to its trend level. Figure 1 depicts the detrended path of monthly industrial production that we observe around the time of the December 1997 crisis along with the detrended dynamic forecasts from our baseline and extended models for the crisis and immediate postcrisis periods. While not capturing the entire decline in output associated with the crisis, the models that add measures of credit market conditions are able to capture parts of the decline that are almost completely missed in models that consider money alone.

We also note that CCI (IV) and DBR help to predict the decline in industrial production better than the observed spread. Since, as we describe earlier, the DBR may purge the spread of its anticipatory component and may itself reflect increases in the cost of credit intermediation and its effect on small and medium-sized firms more precisely than the corporate–government bond spread, we take this as evidence of a credit channel at work.

Results from Structural Models for Production in Heavy and Light Industry

Even though we find evidence of a nonmonetary channel through which Korea's financial crisis was transmitted to the real economy, we suspect that the effects might be different across industrial sectors. To examine this, we repeat the analysis after segmenting the industrial sector into heavy and light industry, and compare the results.[10]

Hayashi and Inoue (1991), among others, also consider light and heavy industry separately in studying firm-level investment, but this segmentation is not typical of the credit channel literature, where the focus is often on comparing small and medium-sized enterprises (SMEs) against larger enterprises.[11] There, it is presumed that supply-side constraints on bank credit will bind SMEs more tightly than larger firms with alternative and lower-cost sources of credit. Potential problems for comparison arise because SMEs need not map neatly into light industry. In Korea's case, however, our comparison of light versus heavy industry, imposed on us by the availability of data over the full period of our study, seems reasonable. For example, data from the *Report on Mining and Manufacturing* of Korea's National Statistical Office indicate that SMEs accounted for 64.2% of the value added in light industry in 1995 and 72.8% in 1999, whereas for heavy industry, SMEs accounted for only 39.2% of the value added in 1995 and 39.7% in 1999. Moreover, the influential chaebols, which specialize predominantly in heavy industry and are among the largest firms in this category, enjoyed preferential access to bank loans that shielded them from the liquidity shock.[12] Given our observation that light industries include generally smaller firms than heavy industries, we would expect light industry firms, on the whole, to have become more credit constrained during the crisis than firms in heavy industry.

Substituting the production indices for heavy and light industry in place of the aggregate index, we repeat our two-stage estimation procedure and report the results in Table 3. The first columns in each panel show that negative unanticipated shocks to money reduce production for heavy industry, with the *F*-test on the lags of money significant at the 1% level, but that light industry is not affected by money shocks at conventional significance levels. This is probably due to demand-side forces in which borrowers in heavy industry reduce their demand for funds when the money supply falls and interest rates rise. Light industry production, for which supply-side credit factors may figure more prominently, appears much less sensitive to the interest rates that lenders offer.

The situation reverses when we focus on measures of credit conditions in the last three columns of each panel in Table 3. Here, increases in all of the credit variables reduce production in light industry much more emphatically than in heavy industry (i.e., the negative and large coefficients on the credit variables are significant at the 1% level for light industry and only at the 10% level for heavy industry). The results are consistent with light industries being more severely affected by credit crunches than the larger and more capital-intensive heavy industries.

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[10] According to Korean Standard Industrial Classification, heavy industry includes manufacturers of coke, refined petroleum products and nuclear fuel, chemicals, nonmetallic minerals, basic and fabricated metals, machinery and equipment, computers and office machinery, electrical machinery, electronics, communication equipment, medical, precision, and optical instruments, watches and clocks, motor vehicles, trailers, other transport equipment, recycling, and pulp and paper products. Light industries includes manufacturers of food, tobacco, textiles, apparel, leather and leather products, luggage and footwear, wood and wood products, publishing, printing, reproduction of recorded media, rubber and plastic products, and furniture.

[11] We also explore the case of large versus small firms. The National Statistical Office provides industrial production indices for large firms and SMEs from 1995. Our estimations imply that the effects of the credit variables on small companies' production are greater than their effects on large firms in both magnitude and statistical significance. The comparison is less useful for our purposes than that of heavy versus light industry; however, data on SMEs are only available from 1995, leaving fewer than 100 time series observations for analysis, and very few from before the crisis for establishing the baseline coefficients.

[12] Borensztein and Lee (2002) present evidence that the preferential access of chaebol firms to bank credit was reduced in the period after the crisis. Their fundamental advantage over smaller firms, however, was certainly not eliminated by the crisis.

Table 3. Second-Stage Results by Industry

	Dependent Variable: Heavy Industry				Dependent Variable: Light Industry			
	Baseline	SPR	CCI (IV)	DBR	Baseline	SPR	CCI (IV)	DBR
Constant	0.586*	0.613*	0.602*	0.575*	0.396*	0.454*	0.469*	0.433*
	(0.149)	(0.149)	(0.149)	(0.148)	(0.140)	(0.137)	(0.140)	(0.137)
$Y_{t\text{-}1}$	0.686*	0.671*	0.678*	0.673*	0.489*	0.455*	0.455*	0.454*
	(0.069)	(0.069)	(0.069)	(0.069)	(0.064)	(0.064)	(0.064)	(0.064)
$Y_{t\text{-}2}$	0.152*	0.161*	0.159*	0.168*	0.427*	0.450*	0.438*	0.454*
	(0.069)	(0.068)	(0.068)	(0.069)	(0.063)	(0.062)	(0.062)	(0.062)
$(M - M^e)_t$	0.443	0.496	0.496	0.612*	0.128	0.201	0.195	0.357
	(0.273)	(0.273)	(0.274)	(0.291)	(0.196)	(0.192)	(0.194)	(0.205)
$(M - M^e)_{t-1}$	−0.266	−0.200	−0.097	−0.181	−0.086	−0.001	0.129	0.025
	(0.275)	(−0.276)	(0.293)	(0.279)	(−0.196)	(0.193)	(0.206)	(0.195)
$(M - M^e)_{t-2}$	0.760*	0.746*	0.831*	0.787*	0.142	0.128	0.234	0.184
	(0.274)	(0.272)	(0.277)	(0.273)	(0.196)	(0.191)	(0.195)	(0.192)
C_t		−0.409	−1.311	−1.826		−0.515*	−1.709*	−2.491*
		(0.214)	(0.813)	(1.099)		(0.153)	(0.587)	(0.785)
Trend	0.001*	0.001*	0.001*	0.001*	−0.000*	−0.000*	−0.000*	−0.000
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
F-test (*p*-value)								
Money shocks	0.934	1.042	1.230	1.218	0.184	0.243	0.558	0.566
	(0.012)	(0.011)	(0.007)	(0.005)	(0.768)	(0.678)	(0.455)	(0.285)

Y is the log of the industrial production index (seasonally adjusted) for heavy and light industry, respectively. The specifications are otherwise the same as those reported in Table 2. The residuals from each specification appear to be stationary on the basis of both ADF and PP tests and are not reported.

* Represents statistical significance at the 5% level. Standard errors are in parentheses beneath the regression coefficients.

Adding SPR, CCI (IV), and DBR reduces the mean-square forecasting errors in dynamic simulations for heavy industry (Figure 2) by 32.6%, 37.2%, and 37.4%, respectively, compared with the baseline over the period from December 1997 to June 1999. For light industry (Figure 3), adding SPR, CCI (IV), and DBR reduces the mean-square forecasting error by 37.8%, 50.3%, and 52.4%, respectively, compared with the baseline over the period from December 1997 to December 1999. Both time periods again correspond to the start of the crisis through the months when industrial production in light and heavy industries returned to their respective trend levels. These exercises confirm that a combination of monetary and credit market factors are needed to explain a reasonably large part of the output contraction associated with the financial crisis, especially in the case of light industry. Further, CCI (IV) and DBR once again perform better than the observed spread in predicting the decline in industrial production, with the improvement most emphatic for light industry, which is consistent with the decline in industrial production being exacerbated by a flight to quality that affected SMEs more severely than larger firms.

6. Vector Autoregressive Analysis

The two-step structural regression models presented in section 5 involve strong assumptions regarding the identification and exogeneity of unanticipated monetary disturbances. In this section, we avoid these potential limitations by focusing on longer-term

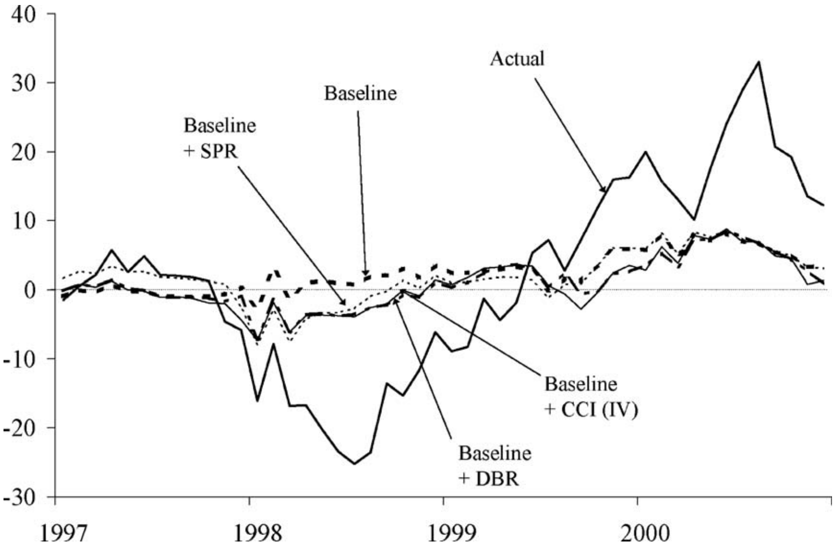


Figure 2. Actual Industrial Production in Heavy Industry and Dynamic Simulations

macroeconomic effects of monetary and credit market disturbances using a series of VAR and LA-VAR models. 2

Baseline VARs

Our baseline VAR models have the following trivariate setup:

$$Y_t = a_{1,0} + \sum_{i=1}^{k} a_{1,i} Y_{t-i} + \sum_{i=1}^{k} b_{1,i} M_{t-i} + \sum_{i=1}^{k} c_{1,i} C_{t-i} + \mu_{1,t}, \qquad (4a)$$

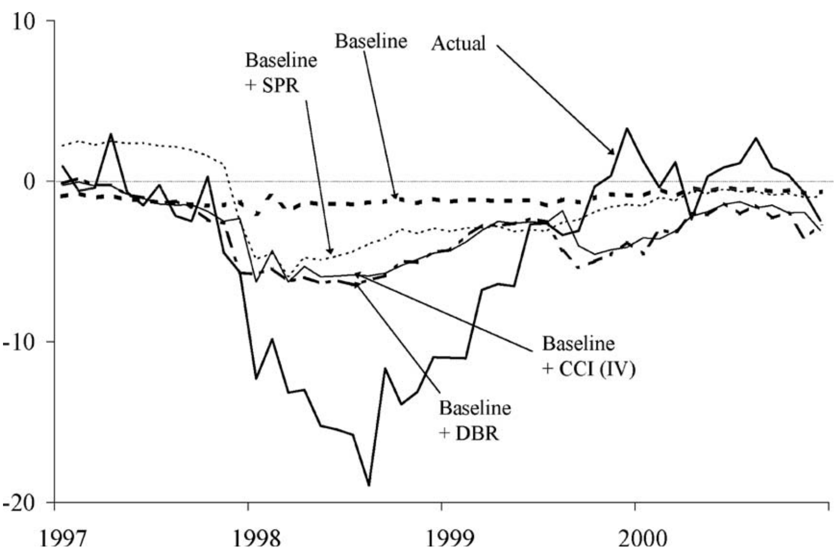


Figure 3. Actual Industrial Production in Light Industry and Dynamic Simulations

Table 4. Unit Root Tests

	Augmented Dickey–Fuller		Phillips–Perron	
	Level	Detrended	Level	Detrended
Log of total real industrial production (*YT*)	−3.192	−3.158*	−4.017*	−4.041*
Log of real heavy industry production (*YH*)	−2.943	−2.905*	−4.356*	−4.380*
Log of real light industry production (*YL*)	−3.178	−2.970*	−3.681*	−3.681*
Log of real money (*M2*)	−2.165	−2.188	−2.061	−2.100
Corporate–Treasury yield spread (SPR)	−4.443*	−4.455*	−4.970*	−4.970*
Dishonored bills ratio (DBR)	−3.692*	−3.707*	−7.586*	−7.586*
Log of real exchange rate (EXR)	−2.450	−2.425	−2.630	−2.633

All variables are in logs except for SPR and DBR. The ADF tests use six lags on the basis of the Akaike information criterion. The level specifications include intercept and trend, whereas the detrended specifications include an intercept only.
* Denotes statistical significance at the 5% level.

$$M_t = a_{2,0} + \sum_{i=1}^{k} a_{2,i} Y_{t-i} + \sum_{i=1}^{k} b_{2,i} M_{t-i} + \sum_{i=1}^{k} c_{2,i} C_{t-i} + u_{2,t}, \qquad (4b)$$

$$C_t = a_{3,0} + \sum_{i=1}^{k} a_{3,i} Y_{t-i} + \sum_{i=1}^{k} b_{3,i} M_{t-i} + \sum_{i=1}^{k} c_{c,i} C_{t-i} + u_{3,t}, \qquad (4c)$$

where k is the lag order, Y is the log of real industrial production, M is the log of real money (*M2*) balances, and C is either the observed corporate–government yield spread or the dishonored bills ratio. The analysis once again runs monthly from January 1987 to December 2003.

Before proceeding to the estimations, it is critical to determine first whether our systems are stationary and thus appropriate for drawing inferences using standard tests of Granger noncausality. To this end, Table 4 reports results from unit root tests for the series used in the

Table 5. Cointegration Tests for VAR Systems with Industrial Production, the Money Stock, and a Measure of Credit Market Conditions

	Trace			Maximum Eigenvalue		
	$r = 0$	$r < 1$	$r < 2$	$R = 0$	$r < 1$	$r < 2$
All industries						
YT-*M2*-SPR	35.24*	13.00	3.28	22.24*	9.72	3.28
YT-*M2*-DBR	26.87	10.94	4.64*	15.53	6.30	4.64*
Heavy industry						
YH-*M2*-SPR	35.44*	12.56	3.59	23.18*	8.67	3.59
YH-*M2*-DBR	32.49*	14.11	4.67*	18.38	9.44	4.67*
Light industry						
YL-*M2*-SPR	30.43*	12.34	3.63	18.08	8.72	3.63
YL-*M2*-DBR	26.36	11.49	4.38*	14.87	7.11	4.38*

Each system includes a measure of credit conditions (SPR or DBR) and the detrended logs of industrial production (*YT*, *YH*, or *YL*) and broadly defined real money (*M2*). Optimal lag lengths were determined for each specification with a series of nested likelihood ratio tests, and each model contains a deterministic drift. The columns labeled $r = 0$ test a null of no cointegration, whereas the $r < 1$ ($r < 2$) columns test a null of at most one (two) cointegrating vectors, with critical values from Table 1 from Osterwald-Lenum (1992, p. 468).
* Represents rejection of the null hypothesis at 5% level.

Table 6. Granger Noncausality Tests for VAR and LA-VAR Systems with the Observed Spread

Sector	Equation	*Y*	*M*	SPR	R^2
All industry	*Y* (4a)	0.821	0.044	−0.380	0.851
		(0.000)	(0.164)	(0.037)	
		[0.000]	[0.177]	[0.023]	
	M (4b)	0.006	0.975	0.115	0.980
		(0.085)	(0.000)	(0.000)	
		[0.061]	[0.000]	[0.000]	
	SPR (4c)	−0.002	0.000	0.764	0.839
		(0.712)	(0.879)	(0.000)	
		[0.599]	[0.907]	[0.000]	
Heavy	*Y* (4a)	0.816	0.019	−0.529	0.814
		(0.000)	(0.063)	(0.071)	
		[0.000]	[0.129]	[0.057]	
	M (4b)	0.011	0.973	0.113	0.981
		(0.031)	(0.000)	(0.000)	
		[0.019]	[0.000]	[0.000]	
	SPR (4c)	−0.003	0.001	0.762	0.839
		(0.656)	(0.916)	(0.000)	
		[0.545]	[0.930]	[0.000]	
Light	*Y* (4a)	0.916	0.023	−0.350	0.890
		(0.000)	(0.447)	(0.061)	
		[0.000]	[0.302]	[0.010]	
	M (4b)	−0.009	0.975	0.048	0.979
		(0.215)	(0.000)	(0.000)	
		[0.245]	[0.000]	[0.000]	
	SPR (4c)	0.001	0.004	0.758	0.837
		(0.553)	(0.757)	(0.000)	
		[0.259]	[0.726]	[0.000]	

The VARs use 10 lags for the systems with total and heavy industrial production, and eight lags for the system with light industry production. Each cell reports the sum of the coefficients for the variable block named in the column label, with *p*-values of *F*-tests for Granger noncausality in parentheses. The *p*-values of *F*-tests for Granger noncausality using the LA-VAR appear in brackets.

VARs that follow, both before and after detrending. The ADF and PP tests reject a unit root for the log of industrial production (total, heavy, and light) after detrending, but do not reject in levels. The spread and the dishonored bills ratio appear to be stationary, while a unit root cannot be rejected for the real money stock.

These tests present some uncertainty as to how the VARs should be specified. If industrial production has a unit root, as the real money stock appears to have, and the system is cointegrated, Sims, Stock, and Watson (1990) show that Wald tests for block exclusion from the VAR in levels are chi-square distributed and thus appropriate for classical hypothesis testing. To examine this possibility, we compute Johansen's (1991) trace and maximum eigenvalue statistics for our trivariate systems and report the findings in Table 5. When the spread is the measure of credit market conditions, both the trace and maximum eigenvalue tests indicate a single cointegrating vector at the 5% level for systems with total industrial production and production in heavy industry, whereas only the trace test confirms a single cointegrating vector for the system with light industry. When the dishonored bills ratio enters

Table 7. Granger Noncausality Tests for VAR and LA-VAR Systems with the Dishonored Bills Ratio

Sector	Equation	*Y*	*M*	DBR	R^2
All industry	*Y* (4a)	0.820	0.027	−1.425	0.814
		(0.000)	(0.121)	(0.010)	
		[0.000]	[0.066]	[0.006]	
	M (4b)	−0.005	0.978	0.370	0.976
		(0.193)	(0.000)	(0.000)	
		[0.286]	[0.000]	[0.000]	
	DBR (4c)	0.004	−0.005	0.839	0.724
		(0.666)	(0.431)	(0.000)	
		[0.612]	[0.321]	[0.000]	
Heavy	*Y* (4a)	0.832	0.006	−2.760	0.736
		(0.000)	(0.003)	(0.126)	
		[0.000]	[0.016]	[0.505]	
	M (4b)	−0.003	0.976	0.456	0.976
		(0.072)	(0.000)	(0.000)	
		[0.075]	[0.000]	[0.000]	
	DBR (4c)	0.004	−0.004	0.778	0.711
		(0.778)	(0.575)	(0.000)	
		[0.880]	[0.405]	[0.000]	
Light	*Y* (4a)	0.901	0.002	−2.271	0.901
		(0.000)	(0.119)	(0.000)	
		[0.000]	[0.279]	[0.000]	
	M (4b)	−0.004	0.971	0.466	0.976
		(0.255)	(0.000)	(0.000)	
		[0.338]	[0.000]	[0.000]	
	DBR (4c)	0.002	−0.003	0.841	0.722
		(0.549)	(0.477)	(0.000)	
		[0.324]	[0.331]	[0.000]	

See note to Table 6. The VARs use 10 lags for the system with all industries, seven for heavy industries, and nine lags for light industries.

as the credit variable, only the trace tests offer evidence of a single cointegrating vector at about the 10% level.

In light of the unit root and cointegration tests, and to ensure that potentially important information in the levels of the data is not lost from overdifferencing, we estimate the VARs in levels both imposing the optimal lag length under cointegration and using the LA-VAR technique proposed by Toda and Yamamoto (1995). The latter show that block exclusion tests still conform to standard Wald distributions if the VARs are estimated with an additional lag and the Granger noncausality tests computed using only the optimally selected lag order. This method, though less efficient when a system is actually cointegrated than constructing Wald tests on the basis of the optimal lag length, is robust to uncertainty about the presence of cointegration and the stationarity of individual variables.

Table 6 shows the VAR results with the observed spread as the indicator of credit conditions. Each row in the table corresponds to a VAR equation as numbered in the text, and the cells include the algebraic sums of the lag coefficients on the variables that appear in the column labels, with the significance level of the *F*-test for Granger noncausality in parentheses. *p*-Values from the corresponding LA-VAR models appear in brackets beneath the Granger

Table 8. Percentage of Variance in Industrial Production Explained by the Credit Variables

	3 months	6 months	12 months	24 months
Ordering: SPR-*M2*-*Y*				
All	10.081	16.949	25.423	25.216
Heavy	6.359	12.391	21.402	21.541
Light	9.683	17.031	23.184	23.009
Ordering: *Y*-*M2*-SPR				
All	2.310	5.648	10.957	11.359
Heavy	1.801	4.951	10.862	11.367
Light	5.074	9.385	13.528	13.165
Ordering: DBR-*M2*-*Y*				
All	4.339	7.057	15.069	15.133
Heavy	1.163	2.243	10.330	10.830
Light	8.950	14.107	30.171	39.351
Ordering: *Y*-*M2*-DBR				
All	1.781	3.465	9.843	9.952
Heavy	0.893	1.633	8.839	9.270
Light	6.024	9.605	23.789	32.546

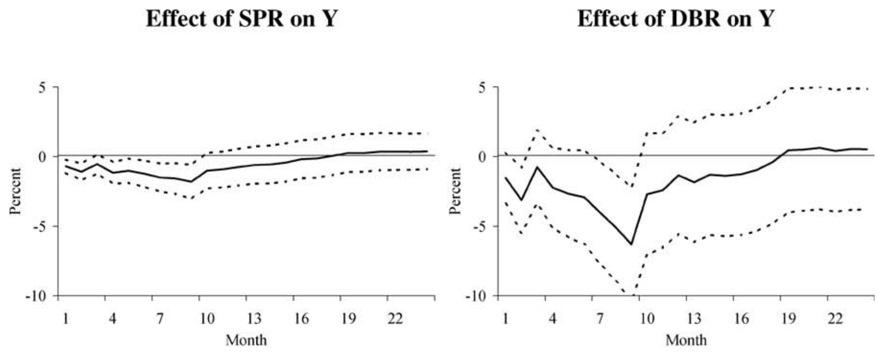
The table shows the percentage of industrial production attributable to the credit variables in our trivariate VAR systems using the optimal lag lengths. For each system, we begin by placing the credit variable first, *M2* second, and industrial production last. We then reverse the ordering in the lower panels.

tests. The upper panel shows results from the VAR for aggregate industrial production, whereas the lower two panels consider production in heavy and light industries separately. In the first system, the observed spread (SPR) Granger-caused industrial production (*Y*) and money (*M*) at the 5% level, but it is not itself Granger-caused by either *Y* or *M*. This result is robust to use of the LA-VAR, and to considering light and heavy industry alone, though in these cases the spread is statistically significant in the *Y* equation at only the 7% level.

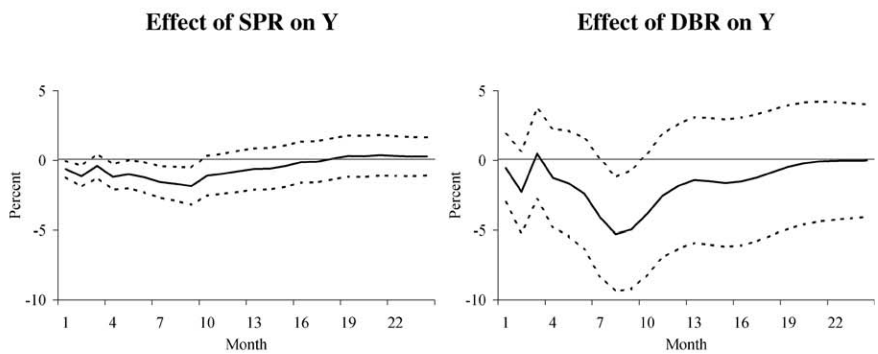
In Table 7, we replace the observed spread with the dishonored bills ratio. The sums of the coefficients on DBR in the output equations (4a) are negative as expected, and DBR is now statistically significant at the 1% level for total and light industrial production in both the VAR and LA-VAR models, but does not appear to affect production in heavy industry. There are no signs of feedback from industrial production or money to DBR in any of the systems.

In Table 8, we report the percentage of the variance in industrial production that can be attributed to the credit variables in our VAR systems at 3-, 6-, 12-, and 24-month horizons. When we order the corporate–government yield spread first, money second, and industrial production last in the upper panel, as suggested by their relative degree of exogeneity in the Granger tests, we find that innovations to the spread explain more than 20% of the variance in all three measures of industrial production after 12 months. When we handicap the credit variable by reversing the ordering in the second panel, we still find that it explains more than 10% of the variance in industrial production.

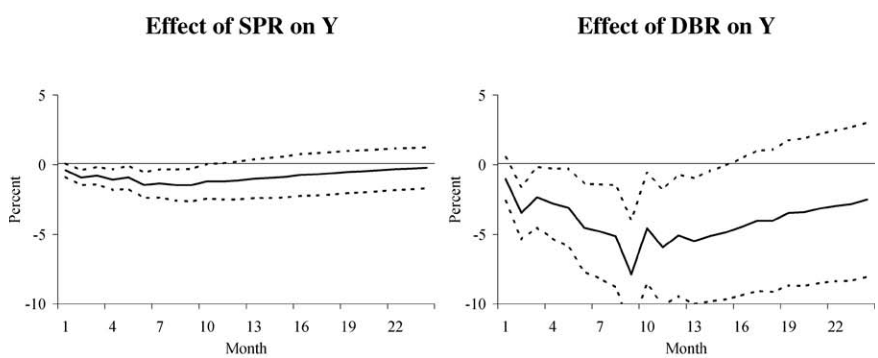
The lower panel of Table 8 uses DBR as the credit variable, and when ordered first, it explains about 15% of the variance in total industrial production, 10% of heavy industry production, and a sizeable 30% of light industry production. The percentages are reduced a bit when the ordering is reduced in the final panel, but less than we observed for the spread. This suggests that the dishonored bills ratio is a particularly strong predictor of industrial production among light industry (i.e., smaller) firms.



(a) all industry



(b) heavy industry



(c) light industry

Figure 4. Impulse Responses of Industrial Production from Trivariate VARs to 1% Innovations in the Credit Variables

Table 9. Granger Noncausality Tests for VAR and LA-VAR Systems with the Observed Spread (and Controlling for the Real Exchange Rate)

Sector	Equation	*Y*	*M*	EXR	SPR	R^2
All industry	*Y*	0.837	0.042	0.014	−0.392	0.791
		(0.000)	(0.022)	(0.634)	(0.123)	
		[0.000]	[0.062]	[0.126]	[0.361]	
	M	0.009	0.977	−0.013	−0.026	0.979
		(0.536)	(0.000)	(0.001)	(0.037)	
		[0.588]	[0.000]	[0.004]	[0.004]	
	EXR	0.015	0.019	0.944	0.291	0.936
		(0.301)	(0.604)	(0.000)	(0.080)	
		[0.293]	[0.623]	[0.000]	[0.139]	
	SPR	0.010	−0.002	0.000	0.783	0.836
		(0.868)	(0.827)	(0.053)	(0.000)	
		[0.783]	[0.910]	[0.051]	[0.000]	
Heavy	*Y*	0.819	0.025	0.023	−0.388	0.739
		(0.000)	(0.026)	(0.691)	(0.198)	
		[0.000]	[0.066]	[0.099]	[0.270]	
	M	0.018	0.973	−0.015	−0.031	0.979
		(0.410)	(0.000)	(0.001)	(0.030)	
		[0.477]	[0.000]	[0.005]	[0.003]	
	EXR	0.004	0.027	0.945	0.272	0.936
		(0.382)	(0.580)	(0.000)	(0.090)	
		[0.411]	[0.652]	[0.000]	[0.157]	
	SPR	0.007	−0.001	0.000	0.782	0.836
		(0.947)	(0.856)	(0.061)	(0.000)	
		[0.929]	[0.932]	[0.059]	[0.000]	
Light	*Y*	0.886	0.019	0.047	−0.270	0.882
		(0.000)	(0.196)	(0.217)	(0.353)	
		[0.000]	[0.359]	[0.216]	[0.597]	
	M	0.012	0.979	−0.018	−0.021	0.978
		(0.767)	(0.000)	(0.002)	(0.029)	
		[0.876]	[0.000]	[0.005]	[0.004]	
	EXR	0.053	0.018	0.920	0.232	0.938
		(0.074)	(0.729)	(0.000)	(0.067)	
		[0.118]	[0.689]	[0.000]	[0.106]	
	SPR	−0.003	0.001	0.003	0.797	0.839
		(0.454)	(0.864)	(0.027)	(0.000)	
		[0.445]	[0.948]	[0.028]	[0.000]	

The VARs use five lags of each variable as determined by nested likelihood ratio tests, and the cells report the algebraic sum of the coefficients on each variable block, with the significance level of the *F*-test for Granger noncausality in parentheses. The *p*-values of *F*-tests for Granger noncausality using LA-VAR are reported in brackets.

We analyze the dynamics of our systems by computing the impulse responses, once again ordering the variables on the basis of their relative exogeneity—SPR or DBR first, *M* second, and *Y* last.[13] The impulse response of aggregate industrial production to 1% shocks in SPR and the DBR are depicted in the left- and right-hand sides, respectively, of panel (a) in Figure 4. The solid lines are the mean impulse responses and the dotted lines are two-standard error bands based on 2500 draws from the posterior distribution of the VAR coefficients. Shocks to

[13] We find that the qualitative results of our analysis of trivariate systems are not sensitive to the ordering of the variables in the VAR.

Table 10. Granger Noncausality Tests for VAR and LA-VAR Systems with the Dishonored Bills Ratio (and Controlling for the Real Exchange Rate)

Sector	Equation	*Y*	*M*	EXR	DBR	R^2
Whole	*Y*	0.843	0.041	0.003	−1.440	0.781
		(0.000)	(0.015)	(0.675)	(0.687)	
		[0.000]	[0.159]	[0.163]	[0.480]	
	M	0.003	0.980	−0.008	0.345	0.978
		(0.613)	(0.000)	(0.000)	(0.126)	
		[0.549]	[0.000]	[0.000]	[0.056]	
	EXR	0.040	−0.001	0.919	−1.800	0.938
		(0.522)	(0.563)	(0.000)	(0.225)	
		[0.469]	[0.482]	[0.000]	[0.140]	
	DBR	0.003	−0.003	0.000	0.822	0.719
		(0.494)	(0.321)	(0.117)	(0.000)	
		[0.540]	[0.073]	[0.019]	[0.000]	
Heavy	*Y*	0.821	0.028	0.008	−1.491	0.726
		(0.000)	(0.017)	(0.731)	(0.797)	
		[0.000]	[0.197]	[0.240]	[0.533]	
	M	0.013	0.975	−0.011	0.214	0.978
		(0.625)	(0.000)	(0.000)	(0.144)	
		[0.537]	[0.000]	[0.000]	[0.050]	
	EXR	0.033	0.004	0.919	−1.877	0.938
		(0.608)	(0.595)	(0.000)	(0.243)	
		[0.517]	[0.498]	[0.000]	[0.148]	
	DBR	0.004	−0.003	0.000	0.801	0.719
		(0.503)	(0.296)	(0.118)	(0.000)	
		[0.688]	[0.065]	[0.018]	[0.000]	
Light	*Y*	0.868	0.015	0.033	−2.330	0.889
		(0.000)	(0.122)	(0.319)	(0.035)	
		[0.000]	[0.334]	[0.076]	[0.214]	
	M	0.006	0.980	−0.010	0.397	0.977
		(0.824)	(0.000)	(0.000)	(0.096)	
		[0.787]	[0.000]	[0.000]	[0.043]	
	EXR	0.079	0.012	0.885	−1.915	0.940
		(0.103)	(0.643)	(0.000)	(0.176)	
		[0.237]	[0.652]	[0.000]	[0.127]	
	DBR	−0.001	−0.001	0.001	0.838	0.718
		(0.546)	(0.410)	(0.123)	(0.000)	
		[0.524]	[0.092]	[0.021]	[0.000]	

See note to Table 9. The VARs use five lags of the system variables based on a series of likelihood ratio tests.

both credit variables have a negative impact on industrial production for at least 10 months (i.e., the point when the two-standard error bands cross the horizontal axis of the response diagrams), bottoming out after seven months and then decaying. Interestingly, the magnitude of the response to shocks in DBR is nearly three times that of the observed spread.

Panels (b) and (c) of Figure 4 show the corresponding impulse responses for heavy and light industry. Here the duration of the effects of shocks to the observed spread and DBR appear to be longer for light industry, lasting for at least 15 months for light industry with DBR as the credit variable and only about nine months for heavy industry.

The results from our trivariate VARs support our view that a flight to quality may have operated during the financial crisis. We say this because the measure of credit conditions that

Table 11. Percentage of Variance in Industrial Production Explained by the Credit Variables (Controlling for the Real Exchange Rate)

	3 months	6 months	12 months	24 months
Ordering: SPR-EXR-*M2*-*Y*				
All	6.101	10.323	18.129	18.108
Heavy	3.685	7.072	13.870	13.795
Light	8.054	12.588	19.191	21.043
Ordering: DBR-EXR-*M2*-*Y*				
All	1.786	4.119	8.297	8.204
Heavy	0.839	2.447	6.099	6.079
Light	5.935	10.469	21.741	28.874

The table shows the percentage of industrial production attributable to the credit variables in our trivariate VAR systems using the optimal lag lengths.

reflects the costs of intermediation for small and medium-sized firms most effectively (i.e., the DBR) outperforms the observed spread in explaining the depth of the crisis. This is especially so for light industry, in which more than 70% of value added is generated by SMEs.

VARs Controlling for the Real Exchange Rate

Even though global factors influence both the real money supply and the credit variables that we have considered in the baseline VARs, these systems do not explicitly control for the role of international shocks in the propagation of the financial crisis. To control for such effects and thereby examine the robustness of our findings, in this section we re-estimate the VAR and LA-VAR models after adding the real effective exchange rate (as report by Korea's National Statistical Office) as a fourth variable.[14]

Table 9 reports the results from Granger noncausality tests for the four-variable systems that use the observed spread as the measure of credit market conditions. They show that both the spread and the real exchange rate Granger-cause money and eventually affect industrial production indirectly through money for all three industrial classifications. Industrial production does not Granger-cause money, the exchange rate, or the spread, except in the case of light industry, where it Granger-causes the exchange rate at the 10% level. The LA-VAR results (*p*-values in brackets) are consistent with these qualitative findings.

Table 10 reports results for four-variable systems with the dishonored bills ratio as the credit variable. In these systems, much of the effect of credit conditions on industrial production also seems to pass through the money stock, though in the system with light industry the direct *F*-test on DBR in the output equation is now also significant at the 5% level.

Table 11 reports the percentage of the variance in industrial production that can be attributed to the credit variables after controlling for the real exchange rate. We place the credit variable first in the ordering, the exchange rate second, money third, and industrial production last, again as suggested by the Granger noncausality tests. In the upper panel, we find that innovations in the observed spread still can explain about 18%, 14%, and 21% of the variance

[14] Johansen tests for four-variable VARs that include industrial production (total, heavy, or light), the money stock, exchange rate, and a credit variable indicating the presence of one (or two) cointegrating vectors in all cases when using the trace statistic, but do not indicate cointegration with the maximum eigenvalue test. This once again leads us to estimate both VARs with the optimal lag lengths under cointegration and the more robust LA-VARs.

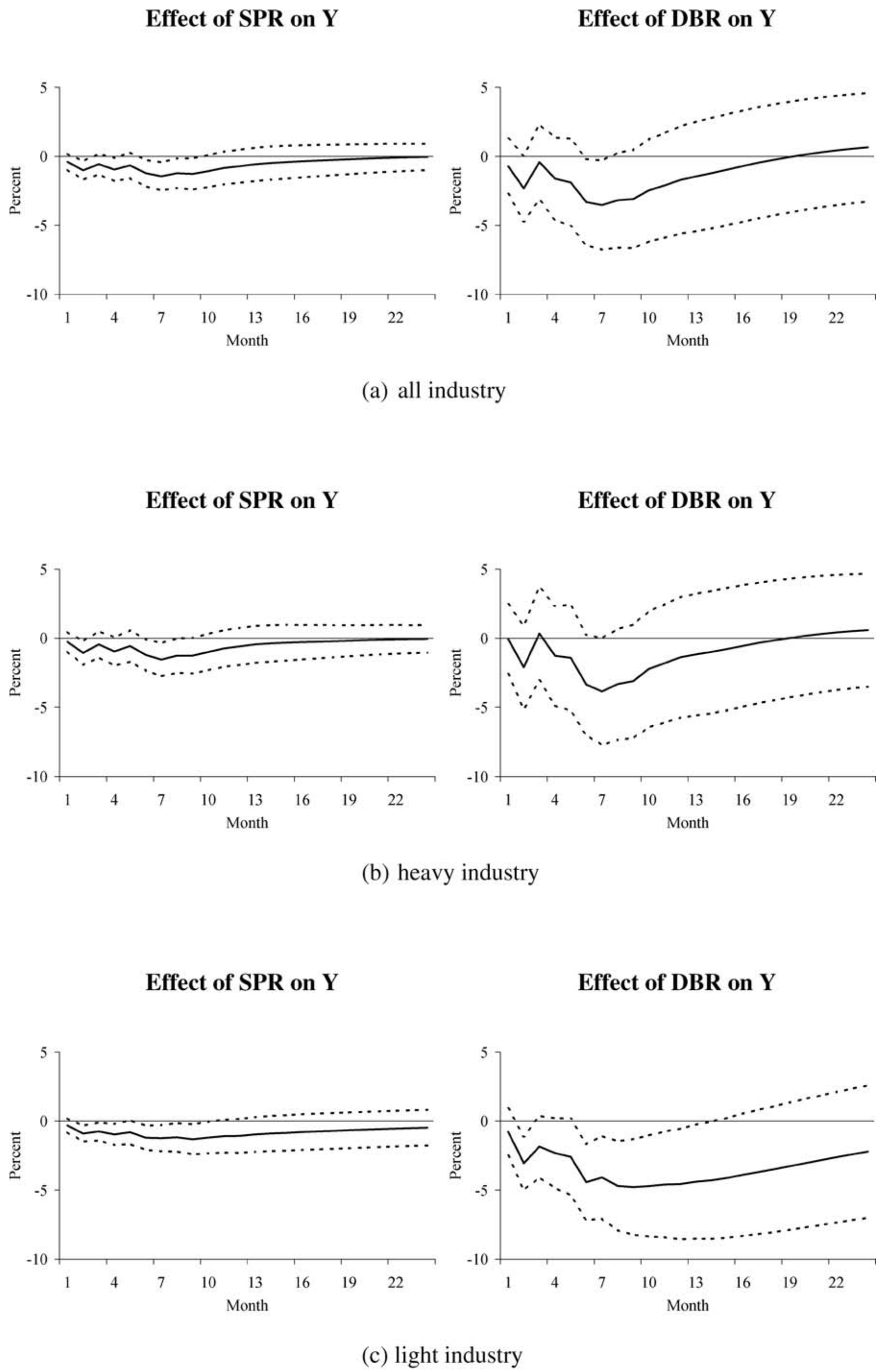


Figure 5. Impulse Responses of Industrial Production in Four Variable VAR Systems to 1% Innovations in the Credit Variables

in total, heavy, and light industrial production, respectively, after 12 months. In the lower panel, DBR explains about 8%, 6%, and 22% of these variances.

Figure 5 shows impulse responses of aggregate industrial production to credit shocks with SPR or DBR placed first in the ordering, followed by the exchange rate, money, and industrial production. Once shocks to the credit variables have worked their way through money, we obtain responses for shocks to the observed spread that are very similar to those in Figure 4 for the three-variable systems. With DBR as the measure of credit conditions, however, we find the magnitude of the responses of industrial production to be even larger that those obtained in the trivariate VARs. Further, in light industry, for which Table 10 shows that shocks to DBR have both direct and indirect effects on production, a 1% shock to DBR lowers *Y* by as much as 6% after 7–9 months. The comparison of the impulse responses for our two measures of credit conditions again shows the DBR to be the better predictor of real activity.

7. Conclusion

Using the ratio of commercial bills dishonored to outstanding bills and the spread between yields on three-year corporate and government bonds as measures of credit conditions, we find empirical evidence that a credit crunch was central to Korea's 1997 financial crisis. The results are robust to whether we use overall industrial production as a measure of real activity or production for heavy and light industries separately, with the effects on light industry clearly the strongest.

We also find that, on the whole, the spread instrumented with the dishonored bills ratio and the DBR itself explain the length and depth of the crisis more completely than the observed interest rate spread, and that this improvement is particularly pronounced for light industry. Since the small and medium-sized firms that account for the majority of light industry output would have had better access to external finance through commercial bills than through the corporate bond market, we attribute at least part of the improved predictive power of the DBR to its ability to capture rising costs of intermediation and an associated flight to quality better than the observed spread.

It thus appears, on the basis of the evidence reported here, that Korea's 1997 financial crisis was driven more by rapid changes in nonmonetary credit market conditions than by monetary shocks, and that the brunt of these changes fell upon firms in light industry, where SMEs dominate. As smaller, innovative firms provided much of the impetus needed to put the nation's postwar "miracle" into motion, policy measures designed to avoid disruptions to the credit allocation process appear to be top priorities for ensuring South Korea's continued financial and economic development.

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