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Business Cycle in Czechoslovakia under Central Planning: Were Credit Shocks Causing It?

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This paper examines credit origins of the business cycle in the former Czechoslovakia. Industrial production is found to be cointegrated with various measures of bank credit during 1976–1990. Noninvestment credits are shown to be Granger-causing industrial production and a feedback relation exists between investment credits and industrial production. Although the impact of credit supply shocks on industrial production has been changing, production decline (growth) seems to follow credit tightening (loosening). However, the paper confirms that credit shocks were only a minor part of the output decline in 1989–1990. *J. Comp. Econom.*, June 1998, **26**(2), pp. 226–245. International Monetary Fund, Washington, DC 20431. © 1998 Academic Press

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1. INTRODUCTION

This paper attempts to provide new evidence on the issue of the credit origins of the business cycle and, ultimately, production decline in the former Czechoslovakia. In the late 1970's and in the 1980's, the Czechoslovak monobank was able to regulate partially the credit supply. Its credit policy changes were not fully or automatically offset by fiscal transfers, changes in prices or arrears. The evidence shows that different measures of credit to the economy and industrial production are cointegrated.

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A relatively strong Granger causality is found to exist between bank credits and industrial production. Noninvestment and total credits are Granger-causing industrial production and a feedback relation exists between investment credits and industrial production. In other words, credit shocks were generating a business cycle. Although the impact of credit supply effects changed during 1976–1990, production decline (or growth) seems to follow credit tightening (or loosening). Our results support the hypothesis that the initial squeeze in credit supply in 1990 (and perhaps beyond 1990) might have contributed to the decline in industrial production. However, the total impact of credit fluctuations was small (immeasurable in a bivariate model) and the production decline was likely generated by and propagated through other mechanisms.

The paper is organized as follows. First, the economic institutions of the former Czechoslovakia are reviewed and links between the real and the monetary economy are outlined. Second, cointegration between bank credits and industrial production is tested. Third, the tests of Granger causality and weak exogeneity for bank credits and industrial production in bivariate vector autoregressions (VAR) are presented. Finally, conclusions are drawn.

2. A NOTE ON THE CREDIT VIEW IN CZECHOSLOVAKIA

2.1. Two Preconditions of the Credit View

The role of credits in the production decline at the outset of economic transition has been debated actively.[2] Calvo and Coricelli (1993) and Calvo and Kumar (1994) noted that, if the supply of credits is cut too abruptly, severe output losses may result.[3] The essence of this hypothesis is that bank lending has only imperfect substitutes and that firms face cash-in-advance constraints to pay for labor and other intermediate inputs. The underlying production function, which we will be using implicitly, is that of Calvo and Kumar (1994).

In order to accept the credit view for a socialist economy, one has to revisit two traditional presumptions concerning command systems. First, credits were endogenous, which meant that the monobank was unable to manipulate the supply of credits, which was production driven. Second, the fiscal system and the direct allocation of inputs fully insulated individual firms from mone-

[2] However, in the former Czechoslovakia, the production decline began with a deceleration of the rate of growth long before the demise of the socialist system in November 1989. Olson (1995) provides an insightful review of socialism's detritus. The output decline in 1990–1993 was most likely overestimated in a fashion similar to that in the other former socialist countries; see Gavrilenkov and Koen (1995).

[3] For a review of the so-called credit view, see Bernanke and Blinder (1988), Bernanke (1993), and Alexander and Caramazza (1994); for the relation between the financial structure and aggregate economic activity, see Gertler (1988).

tary shocks, that is, the assumption of a credit superneutrality. These presumptions effectively exclude credit shocks as a propagation mechanism of the business cycle under central planning.[4]

This paper tries to add credit shocks to the traditional propagation mechanisms of the business cycle. While Bulíř (1995) showed that, in the late 1970's and 1980's, the Czechoslovak monobank was able to regulate the overall credit supply within a monetary targeting framework, this paper revisits the presumption of credit superneutrality. In the period under consideration, and especially from the early 1980's, credit policy changes could not be reversed easily because of fundamental changes in macroeconomic policy design. Although some of the design changes were policymakers' choices, some of them represented the monobank's newly acquired awareness about the inflationary consequences of its actions.

2.2. The Role of Credit in Czechoslovakia

Credit changes are one of the lesser known mechanisms generating business cycles under central planning. While the main reason for the output deceleration in the late 1980's and its collapse in the early stages of the transition can likely be found in the real economy,[5] the credit squeeze might have aggravated further the collapse. Credit was tightened in 1987–1988 and strict ceilings on total commercial bank lending were imposed in 1990 as the State Bank of Czechoslovakia (SBCS) was split in January 1990 into several state-owned commercial banks and the central bank. Those developments led to a vacuum in the credit markets that could not be filled immediately by nonbank institutions and trade credits, given the lack of information, legal framework, and institutions necessary for a private financial market. In addition, the newly created state-owned commercial banks were reluctant to lend in view of their own inexperience in credit risk assessment.

[4] On the issue of a business cycle under central planning, see Bauer (1978), Goldmann and Kouba (1969), Kýn, Schrettl, and Sláma (1978), or Ickes (1990). Recently, the view of a predominantly investment-driven cycle in the former Czechoslovakia has been advocated by Černý and Lazarová (1994) and challenged by Hanousek and Tůma (1996).

[5] The literature highlighted the following real shocks: the simultaneous collapse of the Council for Mutual Economic Assistance (CMEA) and of the Warsaw Pact military procurement, redirecting the use of capital from its value-subtracting planning targets to market utilization, disruptions in oil deliveries from the former Soviet Union, uncertainties related to the formulation of the reform strategy, and major changes in the structure of consumer demand, most notably from domestically produced goods to imports. See Aghevli, Borensztein, and van der Willigen (1992), Borensztein, Demekas, and Ostry (1993), Banerjee (1995), and Fischer, Sahay, and Végh (1996) for empirical analyses.

This paper argues that, even in the prereform period, credit had only imperfect substitutes and its fluctuations contributed to output fluctuations. In contrast to market economies, reinvestment of profits was practically nonexistent, owing to insignificant after-tax profits. Profit taxes were levied at ad hoc rates varying between 70 and 95%. Furthermore, enterprise deposits were largely insensitive to developments in the supply of credits.[6] During 1981–1983, when credit growth declined sharply, those deposits actually increased as a percentage both of credits and of the net material product (NMP). During the second period of credit tightening, in 1987–1988, deposits initially decreased but rebounded in the second year. Moreover, state subsidies to enterprises were cut in the 1980's as well and payment arrears were only imperfect substitutes for bank credits.[7] Assuming that the productivity of most enterprises could not improve immediately after the collapse of the old system, that is, before the full-scale restructuring and privatization, the decline of real credit might have hampered output.

Czechoslovak credit policies in 1976–1990 cannot be explained by the textbook characterization of the command economy that ''money does not matter'' (Grossman, 1990).[8] While large-scale construction and financing of major fixed capital ventures, such as CMEA investment projects and nuclear power plants, were still decided by the planning authority and the monobank had little say in those matters, short-term industrial growth prospects depended on the inflow of new noninvestment and investment capital into the industrial sector.

Following the distinction between noninvestment and investment credits (funds earmarked for certain purposes were not easily interchangeable under central planning), we discuss their roles in output fluctuations. On the one hand, the monobank was subject to quantitative targets on noninvestment credits and decisions about the allocation of those credits were left increasingly in the monobank's hands. On the other hand, the allocations of investment credits and capital expenditure in the budget remained mostly in the

[6] This insensitivity on the aggregate level remains puzzling. It should be noted, however, that enterprise deposits were only about 10% of total enterprise credits throughout the period and that the payment system was based on a simple automaticity: Payments were mechanically credited 10 working days after receiving the invoice. Therefore, firms were forced to maintain large cash positions to avoid an automatic increase in payment arrears.

[7] As in other planned economies, interenterprise arrears were widespread, especially in the 1980's, and served as a sort of trade credit. However, the evidence suggests that arrears entailed some costs for both creditors and debtors. Although Bulíř suggests that changes in payment arrears were credit driven, the buying power of arrears was less than that of credits (Bulíř, 1995, p. 323). Surprisingly, the credit tightening at the onset of the reform was not accompanied by an increase in arrears. The nominal stock of interenterprise arrears declined significantly in the early 1990 and it only began to grow, but stayed well below the 1987–1988 level, during the last quarter of the year.

[8] The Czechoslovak economy did resemble the textbook case in 1950–1965, however.

hands of the planning authority, although the importance of centrally allocated investment financing declined over time.[9]

In this paper, we stress the role of noninvestment credits in the propagation mechanism of the business cycle. Those credits were used for financing inventories, wages, and intermediate production, for which rudimentary markets were in place. In other words, those credits were mainly servicing short-term needs of the supply side of the economy. From the early 1980's, the granting of noninvestment credits was at the discretion of the monobank; it ceased to be automatic or tied to material flows. Monobank branches executed a sort of project assessments for noninvestment credits, for example, by calculating inventories turnover and return-on-assets ratios. Bargaining on the side of firms was usually required to obtain the demanded volume of credits. However, there was no (or very limited) rationing by interest rates even though the effective interest rates rose steadily from the late 1970's.

Investment credits were used only for procurement of fixed capital and were generally at the discretion of the planning authority, that is, the portion of the investment contained in the plan was essentially financed automatically by the monobank.[10] To be precise, the monobank usually financed the gap after the firms' own resources and capital expenditures by the state budget had been apportioned. However, from the early 1980's, industrial firms increasingly gained freedom in determining the volume and structure of their fixed capital investment and in determining the sources of its financing. For example, investment projects totaling Kčs 3 million and eventually Kčs 10 million (about \$0.3 million and \$1 million at the official exchange rate) were not subject to planners' approval. The share of those small investment ventures was gradually increasing; for example, between 1985 and 1988 the value of the former in current prices increased by 70%, while centralized investment stagnated (Federální ministerstvo financí 1990, p. 178).

Our version of the credit view stresses the importance of noninvestment credit. Hence, we expect changes in noninvestment credits to precede changes in industrial production, while industrial production expansion would drive investment credits. Moreover, we expect to find feedback relationships be-

[9] The share of decentralized financing was increasing fast. While only 44% of firms' investment was financed from own financial resources in 1976–1980, this share increased to 64% in 1986–1989 (Federální ministerstvo financí, 1990, p. 175). During the same period, the share of credits decreased marginally from 28 to 26% and the share of the state budget dropped from 28 to 10%. In the 1950's and early 1960's, 45–55% of investment was budget financed while the rest came from firms' financial resources. Although investment credits were introduced in 1959, their role remained marginal until 1967.

[10] It is well known that the objective function of central planners embraced technologies that simply maximized the growth of NMP (Easterly and Fischer, 1994). Resources under central control were directed toward capital-intensive projects and project selection was biased in favor of large or already producing firms.

tween investment credits and output that capture the capital-intensive character of the economic growth in the socialist economies. Those hypotheses are tested in the next section using Granger causality tests.[11]

Although Granger causality tests cannot prove a functional relationship, the existence of a third variable driving credits and industrial production simultaneously would be hard to justify. The planning authority lacked both the information and the instruments necessary for control over the allocation of noninvestment resources; its financial control over investment was also eroded over time.[12] It is notable that most of the noninvestment and investment inputs were available without significant queues and little rationing was enforced. In contrast to other socialist economies, the input markets were not in a state of permanent global shortage, albeit local and temporary disequilibria developed occasionally (see Dlouhý, 1988 and Klaus, 1990).

2.3. *Credit Targeting*

In this section, we describe briefly the mechanism that guided Czechoslovak credit policies on the macrolevel. From the early 1970's, a simple rule was sought for a monetary policy that would shield the monobank from firms' demands for additional credits, especially those financing working capital. The basic monetary target (BMT; *základní monetární kritérium*) was defined in 1976 as a relationship between the growth rate of noninvestment credit to the enterprise sector and the growth rate of net material product.[13] Although the BMT was not defined in any law, it was outlined in monobank's internal documents (e.g., Státní banka československá, 1981). The annual targets (and five-year indicative benchmarks) were negotiated between the monobank and the planning authority and eventually decreed in the Economic Memorandum of the Government. In principle, the monobank could have been held accountable for any overruns, although this was never the case.

One can rewrite the BMT as

$$U^*_{t+1} = U_t(1 + r^*),$$

[11] There is an extensive literature using VAR models for testing the money-versus-income causality in developed countries. In addition to the pioneering paper by Sims (1972), see, for example, Friedman and Kuttner (1992) and Becketti and Morris (1992).

[12] Some authors even argued that, in the mid-1980's, the Czechoslovak planning authority lost its ability to control aggregate resource allocation (Hlaváček, 1990; Klaus and Tříska, 1989; and Mlčoch, 1990). In their view, economic policy was conducted as a three-sided cooperative game, with the players as the firms, the planning authority, and the monobank. Although none of the players had a dominant role in the decision process, there was no central planner in the original sense of the word and the firms had information superiority.

[13] This pragmatic, quasimonetarist criterion gained support among certain Czech economists (Klaus and Rudlovčák, 1979 and Kočárník, 1983).

where U^*_{t+1} is the targeted stock of noninvestment credits on December 31 in year $t + 1$; U_t is the actual stock of noninvestment credits on December 31 in year t; and r^* is a discretionary growth coefficient based on an implicit feedback function embodying the state of the economy.[14]

Although the BMT had several design flaws, most notably it failed to control credit expansion outside the target point of $t + 1$, it played an important role in constraining the demand for noninvestment credits. The political economy of the feedback function and some empirical analysis of the policies associated with the BMT are discussed in Bulíř (1995).

3. IS THERE A LONG-TERM RELATIONSHIP BETWEEN CREDITS AND PRODUCTION?

3.1. Data and Time Series Properties

Noninvestment and investment credits along with industrial production at quarterly frequencies are utilized in this study (Fig. 1).[15] The sample periods are the first quarter of 1976 to the fourth quarter of 1990, as dictated by the availability of the original data taken from the SBCS's database and from *Statistické obzory,* a monthly publication of the Federal Statistical Office. Since both time series are in current prices, there is no need to deflate either variable.

Total credit data consist of two time series. About two thirds of the total credits to the economy were noninvestment credits and one third were investment credits. Given the accounting practices, noninvestment credits could not be used to finance investment projects and vice versa. Moreover, firms could neither borrow abroad nor borrow outside the monobank. The latter condition was, however, violated occasionally by the existence of interenterprise arrears.

Industrial production data were collected for all centrally planned enterprises, which constituted more than 95% of the industrial base. The original

[14] The monobank and the government used several definitions for the BMT: non-investment credits only, total credits to the economy, or total credits excluding the foreign trade credits and the investment credits for the CMEA projects. Because the monobank's control over foreign trade credits and certain types of investment credits was limited, the last definition was the operational measure used by the SBCS's staff. It was precisely this definition that was used by the former chairman of the SBCS to praise the monobank, ''. . . [in 1981–1985] the credit growth was lower than that of the nominal NMP,'' see Stejskal (1986, p. 75). Next year he specified, ''. . . this year we want the growth rate of credits to be lower by 1.3 percentage points than the growth rate of the net material product,'' see Stejskal (1987, p. 77). In the mid-1980's, the BMT was augmented by several microeconomic criteria to create the so-called criteria of credit efficiency (Kroupar, 1987).

[15] Unfortunately, appropriate quarterly data for other variables affecting the business cycle, e.g., labor force and capital stock, fiscal subsidies and taxation, an effective exchange rate, terms of trade, or foreign loans, are not available. Hence, a multivariate VAR that would gauge the relative contribution of credit shocks could not have been performed.

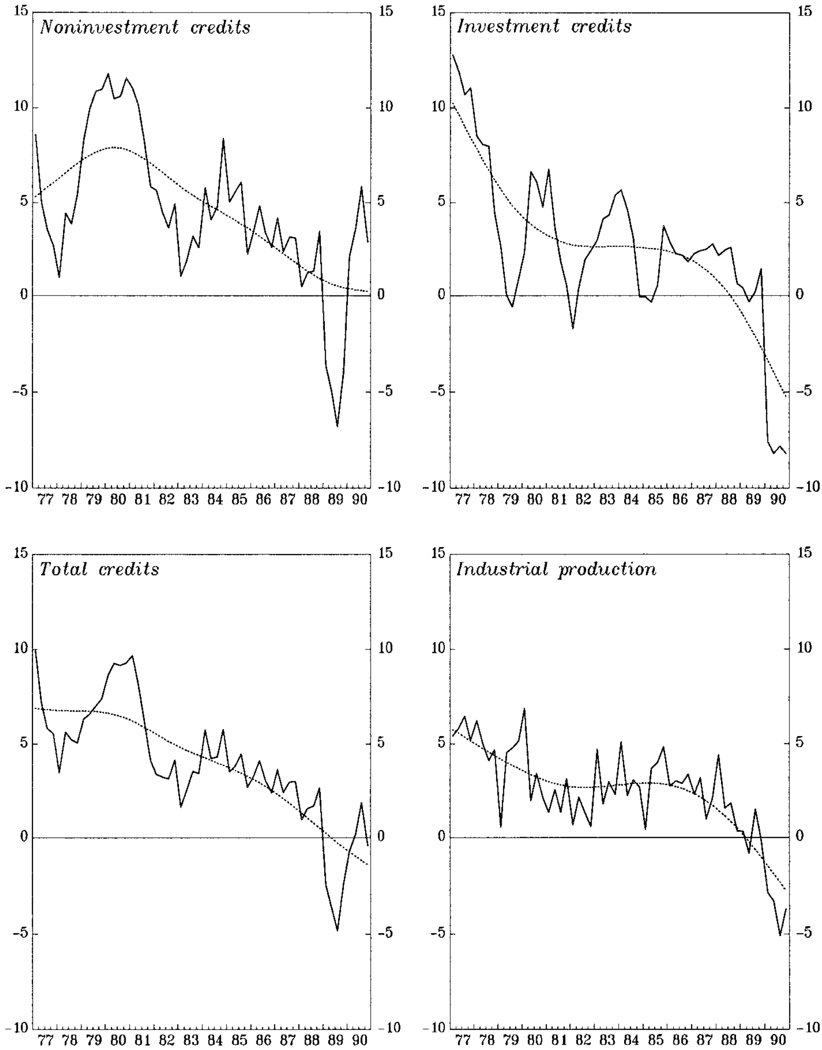

FIG. 1. Growth in credits and industrial production (seasonally unadjusted year-on-year percentage change and Hodrick–Prescott filter). *Source.* State Bank of Czechoslovakia, Federal Statistical Office.

time series were published as percentage change over the same period of the previous year (calculated from current prices); we recovered the nominal values from the absolute industrial production data for 1989, which were published in 1991.

TABLE 1

Dickey–Fuller and Augmented Dickey–Fuller Tests for the Presence of a Unit Root, 1976:I–1990:IV

Variable[a]		DF[b]	ADF(1)[c]	ADF(2)	ADF(3)	ADF(4)
Credits						
Noninvestment	[d]	−1.505	−1.119	−0.461	0.578	−0.954
	[e]	−7.157	−4.749	−3.765	−4.943	−3.574
Investment	[d]	−3.746	−2.627	−1.881	−1.129	−1.171
	[e]	−8.811	−6.652	−4.188	−5.678	−5.510
Total	[d]	−0.791	−0.313	0.240	1.314	0.399
	[e]	−10.643	−6.255	−3.437	−3.080	−3.017
Industrial production	[d]	−8.745	0.012	−0.464	1.037	0.736
	[e]	−12.485	−6.944	−4.735	−5.624	−4.160
Critical values[f]	[d]	−3.486	−3.488	−3.489	−3.490	−3.492
	[e]	−3.494	−3.495	−3.497	−3.499	−3.501

[a] All variables are in natural logarithms.

[b] DF is the Dickey–Fuller Statistics. The regression equation contains a constant and a trend.

[c] ADF(k) is the Augmented Dickey–Fuller Statistics with lag k. The regression equation contains a constant and linear trend.

[d] Levels.

[e] First seasonal differences.

[f] At the 95% confidence interval.

Observing that all variables are nonstationary in levels is easy as documented by the Dickey–Fuller and the Augmented Dickey–Fuller tests in Table 1. However, after taking successively seasonal differences and first differences, all time series begin to exhibit mean-reverting properties, which confirms that the original series were seasonally integrated of the order one (SI(1)).

Empirical analyses around points of structural or institutional breaks, such as those in the mid-1980's and the early 1990's, raise the question of analytical consistency. We are convinced, however, that from the institutional point of view the 1990 economy is not significantly different from the 1986 economy, for example. Prices were liberalized in January 1991 only; privatization started in 1992 only. Of course, the same claim would be harder to justify for 1991 or even 1992. It is primarily for this reason that our analysis ends in 1990.

3.2. Cointegration

Various measures of credit and industrial production moved together during the sample periods. Table 2 provides Johansen's tests (JJ) of the cointegrating relationship between industrial production and various definitions of credit

TABLE 2

Johansen Test Statistics for Cointegration, 1976:I–1990:IV (Vector Autoregression with Two Lags, Unrestricted Intercepts, and No Trend)[a]

	Maximum eigenvalue test statistics	
Hypothesis:[b]		
Null	$r = 0$	$r \leq 1$
Alternative	$r = 1$	$r = 2$
Industrial production and		
Noninvestment credits	17.24	9.32
Investment credits	22.53	9.91
Total credits	19.92	6.35
Critical values[c]	14.88	8.07
	Trace test statistics	
Hypothesis:[b]		
Null	$r = 0$	$r \leq 1$
Alternative	$r = 1$	$r = 2$
Industrial production and		
Noninvestment credits	26.56	9.32
Investment credits	32.45	9.91
Total credits	26.27	6.34
Critical values[c]	17.86	8.07

[a] The lag structure selection was based on three tests: an iterative application of restrictions on a higher order VAR [Holden and Perman (1994)], the Akaike information criterion, and the adjusted coefficient of determination.

[b] The value for ‘‘r’’ is the number of cointegrating vectors.

[c] At the 95% confidence interval.

variables.[16] The null hypothesis of no cointegrating vectors can be rejected in favor of the existence of two cointegrating vectors for all equations but those including total credits, implying Granger causality in at least one direction. As supplemental evidence, we also performed the Engle–Granger (EG) test of cointegration with mixed results. The Durbin–Watson tests (DW) infer cointegration, but the augmented Dickey–Fuller tests, with the exception of

[16] The relevant references for the cointegration techniques are Johansen and Juselius (1990), Engle and Granger (1987), and Urbain (1992). The usual lag-selection tests, i.e., the Akaike information criterion and the adjusted R^2, were used. The long-run matrix π in the JJ procedure can be decomposed as $\alpha\beta'$, where α and β are $kx1$ vectors. We normalize the long-run coefficients β' as $(1, -\delta')$ and partition the adjustment matrix as $\alpha' = (\alpha_1, \alpha_2)$. It should be noted that the first rows of the estimated adjustment matrices (α_1) in the JJ procedure are negative, which is consistent with the hypothesis of an error correction mechanism.

TABLE 3

Cointegrating Vectors from the Johansen and Engle–Granger Procedures

	Johansen estimation β_{JJ}[b]		Engle-Granger estimation			
Variables[a]	First vector	Second vector	β_{EG}[c]	R^2	DW	ADF(1)[d]
Noninvestment credits	0.384	0.636	0.521	0.87	2.56	−2.73
Investment credits	1.052	0.436	0.943	0.90	1.96	−3.44
Total credits	0.351	0.677	0.622	0.90	2.79	−2.41
Industrial production[e]	0.951	2.296	0.956	0.90	1.82	−4.37

[a] Right-hand side variables in the regression of industrial production on credit variables. For the sake of simplicity, only those cointegrating relationships are presented for which Granger causality was later identified.

[b] Normalized parameters of the long-term relationships from the Johansen procedure (JJ).

[c] A parameter of the long-term relationships from the Engle–Granger regressions (EG).

[d] The critical value at the 95% confidence interval is −3.44.

[e] Industrial production regressed on investment credits.

the regression of investment credits on industrial production, suggest otherwise (Table 3). The latter result can be attributed, however, to the generally low power of the EG test rather than to the lack of cointegration: Multiplicity of cointegrating vectors in the JJ test generally signals a nonrobustness of the EG test.

Even though credit inflows had obviously no multiplicative effects on production, the estimated long-run elasticities are sensible and clearly reject the hypothesis of credit neutrality (Table 3). These estimates also do not support the hypothesis of outright wasteful and inflationary credit decisions under central planning as described for the former Soviet Union by Easterly and Fischer (1994). Noninvestment and total credit elasticities of industrial production are in the range 0.4–1.0; industrial production elasticity of investment credit is estimated to be between 1.0 and 2.3. Moreover, those estimates are generally consistent across different estimation techniques.

4. GRANGER CAUSALITY RESULTS

4.1. Directions of Granger Causality

The results support the hypothesis that noninvestment credits led industrial production during 1985–1990 and that investment credits had a feedback relation with production during the same period. Therefore, assuming that productivity of most Czechoslovak enterprises could not improve immediately

TABLE 4

Direction of Granger Causality[a]

Variable	1976–1984	1985–1990	1976–1990	Variable
Noninvestment credits	—	→	→	Industrial production
Investment credits	←	←, ↔	←, ↔	Industrial production
Total credits	—	→	→	Industrial production

Source. Tables 5 and 6.

[a] A ''—'' signals that no Granger causality was detected. A ''→'' (''←'') signals that the left-hand (right-hand) side variable is Granger-causing the right-hand (left-hand) side variable. A ''↔'' signals a feedback relationship between the left-hand and right-hand side variables.

after 1989 and before full-scale restructuring and privatization, lower credit supply (or growth) in 1989–1990 might have adversely affected output (or growth) at the outset of the reform. We will argue, however, that a quantitative answer to the credit-squeeze-production-decline puzzle cannot be obtained from bivariate VARs.

How stable was the Granger causality over time? Earlier we discussed literature suggesting that the partially independent position of the Czechoslovak monobank developed over time as its intermediation powers were increasing. Mechanical tests of statistical significance placed the structural break either in 1984 or in 1985. Hence, we estimated our equations for three periods: the full sample (1976:I–1990:IV), the preindependence period (1976:I–1984:IV), and the independence period (1985:I–1990:IV).

The regression estimates reveal directions of Granger causality conforming with the assumptions outlined earlier. See Table 4 for a summary and Table 5 and 6 for the respective marginal significance levels of the likelihood ratio tests.[17] First, the VAR(2) models show that total credits and noninvestment credits were Granger-causing industrial production in 1985–1990. Therefore, new credits financing inputs (mainly labor and inventories) were preceding changes in industrial output. However, noninvestment and total credits do not seem to be Granger-causing industrial output in 1976–1984. Hence, as expected, credits may have been neutral in the period prior to the mid-1980's. Industrial production appears to be Granger-causing neither noninvestment nor total credits in any period.

[17] Note that we employed two complementary definitions of Granger causality tests, each with two differently transformed time series (see Appendix). As the tests of Granger causality have generally low power in small samples, it does not come as a surprise that one out of four tests is usually a substantial outlier. As a result, we consider the 25% significance level a reasonable benchmark for our purposes.

Second, industrial production was Granger-causing investment credits in both periods. Moreover, the tests also suggest the possibility that investment credits Granger-cause industrial production. Hence, a feedback relation exists between investment credits and industrial production in the second period. While the former finding conforms with the usual notion of capital-intensive growth under central planning, that is, industrial production growth generated its own demand for investment credits, which would finance further expansion of fixed capital, the latter finding hints at the credit view.

4.2. Exogeneity Tests

Even though we have established Granger causality, can we be sure that credit supply was set independently from output developments? In our case, industrial production is postulated to be explained by noninvestment credit fluctuations during certain periods, while in other periods credit fluctuations are thought to be driven by production changes. To overcome the uncertainty about the exogenous vs endogenous character of individual variables, we tested for weak exogeneity as a supplement to Granger causality in the cointegrating regression described in Table 2.[18]

Those results complement our Granger causality findings, namely the link from noninvestment and total credits to industrial production and from industrial production to investment credits. Noninvestment and total credits are found to be weakly exogenous with respect to industrial production (the likelihood-ratio tests yield 0.274 and 3.088, respectively) and industrial production is found to be weakly exogenous with respect to investment credits (0.462). Therefore, given the earlier established direction of Granger causality for the variables in question, noninvestment and total credits are found to be strongly exogenous with respect to industrial production and industrial production is found to be strongly exogenous with respect to investment

[18] In this section we are drawing on Urbain (1992), that is, on exogeneity modeling. The tests used are those suggested by Deadman and Charemza (1992). A variable y_t is said to be weakly exogenous with respect to x_t for the parameters of interest ψ, if knowledge of ψ is not required for the inference on the marginal process of y_t. In other words, weak exogeneity implies that a single-equation regression can be estimated efficiently. The formal test is as follows. If the second rows of the adjustment matrix in the JJ procedure contain zeros ($\alpha_2 = 0$), then the variable in question is said to be weakly exogenous with respect to the long-run parameters. The likelihood-ratio test for this hypothesis is distributed as $\chi^2(1)$ under the null hypothesis with a critical value of 3.841 at the 5% significance level. Finally, once a variable y_t is found to be weakly exogenous with respect to x_t and past observations of x_t are not a Granger cause for y_t, y_t is considered to be strongly exogenous with respect to x_t for the parameters of interest ψ. In turn, strong exogeneity implies a single-equation forecasting model.

TABLE 5

Significance Levels of Granger Causality Tests: Production Causes Credits[a]

	Noninvestment credits	Investment credits	Total credits
	Granger approach Deterministic trend and seasonal dummies		
1976:1–1984:4	29	1	36
1976:1–1990:4	66	1	99
1985:1–1990:4	80	3	96
	Without deterministic variables		
1976:1–1984:4	86	1	71
1976:1–1990:4	86	59	68
1985:1–1990:4	85	39	19
	Sims approach Deterministic trend and seasonal dummies		
1976:1–1984:4	69	1	18
1976:1–1990:4	32	1	53
1985:1–1990:4	47	0	11
	Without deterministic variables		
1976:1–1984:4	49	4	54
1976:1–1990:4	79	21	95
1985:1–1990:4	90	3	49

[a] The numbers in each column are marginal significance levels for the likelihood-ratio test of the joint hypothesis that all of the estimated coefficients of industrial production are equal to zero. For example, the ''29'' in the first row and column indicates that those coefficients are statistically different from zero at the 29% significance level.

credits. Production, however, fails the weak exogeneity test with respect to noninvestment and total credits (7.856 and 10.064, respectively) as do investment credits with respect to industrial production (8.045).

4.3. Contribution of Credit Shocks to the Business Cycle

The VAR estimates also allow us to construct impulse-response functions, which measure the quantitative impact of a unitary change in credit variables on industrial production and of a unitary change in industrial production on investment credits (Fig. 2). The values of the estimated impulse response

TABLE 6

Significance Levels of Granger Causality Tests: Credits Cause Production[a]

	Noninvestment credits	Investment credits	Total credits
	Granger approach Deterministic trend and seasonal dummies		
1976:1–1984:4	89	18	94
1976:1–1990:4	8	14	2
1985:1–1990:4	21	4	5
	Without deterministic variables		
1976:1–1984:4	38	27	82
1976:1–1990:4	59	23	69
1985:1–1990:4	22	31	21
	Sims approach Deterministic trend and seasonal dummies		
1976:1–1984:4	72	59	83
1976:1–1990:4	29	3	18
1985:1–1990:4	11	1	12
	Without deterministic variables		
1976:1–1984:4	29	75	28
1976:1–1990:4	97	32	68
1985:1–1990:4	22	39	59

[a] The numbers in each column are marginal significance levels for the likelihood-ratio test of the joint hypothesis that all of the estimated coefficients of credit variables are equal to zero. For example, the ''89'' in the first row and column indicates that those coefficients are statistically different from zero at the 89% significance level.

functions are relatively large. A one-time temporary increase in total or noninvestment credits by 1% increases industrial production by a cumulative 0.9–1.1% over six quarters, by which time production function has returned to its baseline. Most of this effect happens in the first two quarters. This finding is consistent both with the credit view, which assumes that the credit shocks that affect the working capital should be strong but short-lived, and with our earlier estimates of the long-run credit elasticities of industrial production in cointegrating regressions.

Although the VAR estimations reported above seem to be robust and stable, their predictive power declines during the late 1980's. Most notably, the VAR models fail to predict the production slowdown in 1988–1989 and its fall in

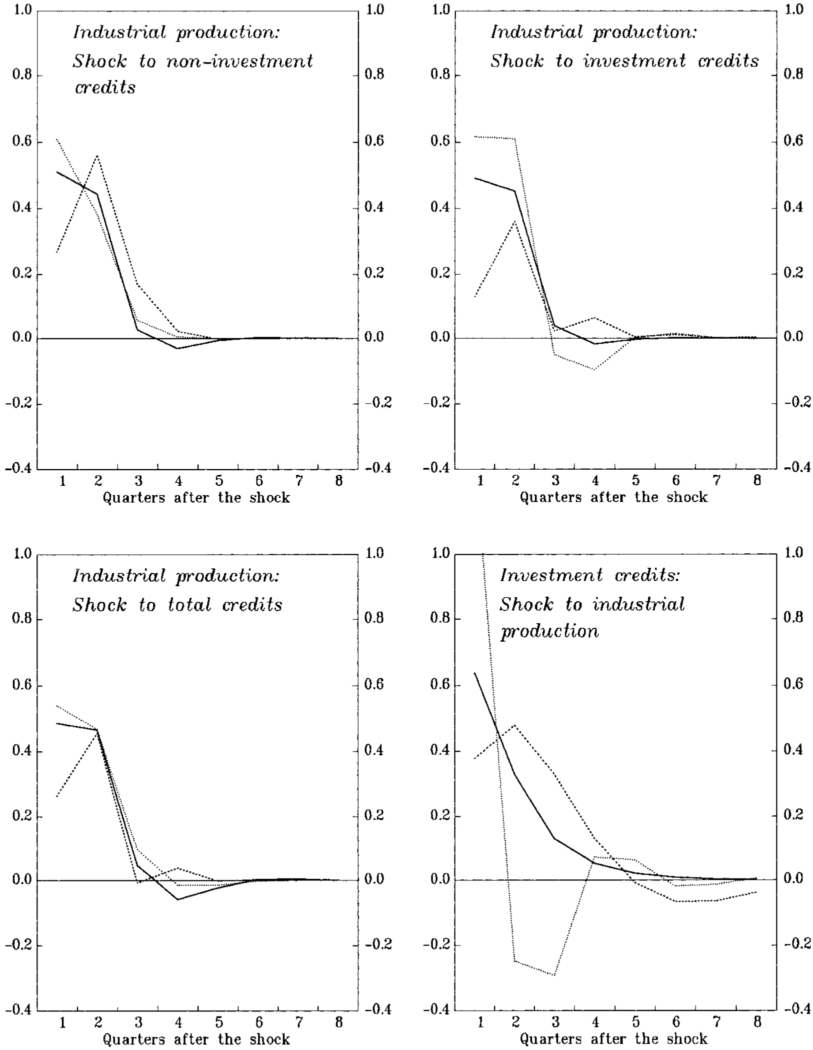

FIG. 2. Impulse response functions (percentage change, relative to baseline, for a 1% temporary shock to the variable indicated). *Source.* Author's calculations.

1990 in both one-step ahead and multistep forecasts. This can be attributed to two features of the simple bivariate VAR models. First, the reduced form models omit shocks to several relevant variables, e.g., fiscal policy, labor inputs, exchange rates, changes in the export–import regime, and political instability, for which reliable data are not available although they clearly had an impact on industrial production.[19] In a multivariate model, those shocks would have to be accommodated by a looser credit policy if the industrial output were to stay unchanged. Second, the negative credit shock was simply too small relative to other shocks to account for a visible part of the industrial production decline in 1990.

5. CONCLUDING REMARKS

This paper has evaluated quantitatively the hypothesis that a credit squeeze might have contributed to the output decline in the former Czechoslovakia. The institutional setup of the Czechoslovak economy in the 1970's and 1980's hints at the monobank's ability to regulate the supply of noninvestment credits, even though its control over investment credits was limited. The results from the cointegration tests suggest existence of a long-run relationship between the real and monetary sectors of a planned economy.

Credit shocks apparently explain a part of the business cycle in the prereform Czechoslovakia. Three main inferences stand out. First, industrial production Granger-causes investment credits, but industrial production does not Granger-cause noninvestment or total credits. Second, noninvestment and total credits Granger-cause industrial output during the period 1985–1990, but not before. Third, a feedback function might exist between production and investment credits during the period 1985–1990. Moreover, the results of Granger causality tests are complemented by tests of weak exogeneity.

The results seem to support indirectly the hypothesis that the credit squeeze at the outset of the Czechoslovak reform (and perhaps beyond 1990) contributed to the decline in industrial production as monetary policy did not accommodate fully the adverse shocks of 1989–1990. The estimated impulse response functions suggest that the fall in industrial output would follow quickly after the credit squeeze with a lag of two quarters at maximum. There is little evidence from our models, however, that the credit shock was a major cause of the output decline. To quantify its impact, one would have to employ multivariate VARs.

APPENDIX: TESTS OF GRANGER CAUSALITY

The following two tests were used in the analysis above. It should be noted that the Granger and Sims tests are not considered to be substitutes but rather complements; see Charemza and Deadman (1992). In other words, there is no tradeoff between these two sets of results and a rule of thumb must be

[19] The very fact that two cointegrating vectors are found might suggest that the endogenous versus exogenous division of variables is imperfect and that the true production function should consist of not one but two or more equations (Charemza and Deadman, 1992).

used to evaluate them. In both cases, we are testing whether x is Granger-causing y using VAR(2) models.

1. *Granger approach.* Estimate

$$y_t = A_0D_t + \sum_{j=1}^{2} \alpha_j y_{t-j} + \sum_{j=1}^{2} \beta_j x_{t-j} + \epsilon_t;$$

if $\beta_1 = \beta_2 = 0$, then x does not Granger-cause y. If the null hypothesis is rejected then x Granger-causes y. This is clearly a straightforward test of variable deletion.

2. *Sims approach.* Estimate

$$x_t = A_0D_t + \sum_{j=1}^{2} \gamma_j x_{t-j} + \sum_{j=-2}^{2} \delta_j y_{t-j} + \upsilon_t;$$

if $\delta_{-1} = \delta_{-2} = 0$, then x does not Granger-cause y.[20] If the null hypothesis is rejected then y does not cause x; that is, x Granger-causes y. This type of test is less straightforward, since one is assuming that the future cannot cause the present and, hence, future values of y cannot cause the current values of x. Indeed, a logical conclusion of finding nonzero coefficients on leading y terms is that x is a Granger-cause for y.

We estimate two sets of equations, both based on the above VAR models. The first set has a deterministic time trend and a deterministic seasonal factor. All variables are in levels and the equations include the following deterministic variables: an intercept, linear time trend, and three seasonal dummies. The introduction of a time trend and seasonal dummies is expected to alleviate the problem of nonstationarity in variables expressed in levels. The second set is without a time trend and is deseasonalized. The equations include only one deterministic explanatory variable, an intercept. All variables are in first seasonal differences ($\Delta x_t = x_t - x_{t-4}$), which were subsequently subject to first difference ($d\Delta x_t = \Delta x_t - \Delta x_{t-1}$). The twice-differenced variables are stationary, as demonstrated by the Augmented Dickey–Fuller tests, reported in Table 1.

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[20] Note that $\delta_{-1}, \ldots, \delta_{-2}$ are parameters of lead variables.

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