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# **An Analysis of Repressed Inflation in Three Transitional Economies**

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**An assessment of the extent of repressed inflation in  
Czechoslovakia, Poland, and Romania between 1980 and 1990.**

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In centrally planned economies, domestic prices do not respond flexibly to market forces, so economic disequilibria — including repressed inflation — persist. Feltenstein and Ha assess the extent of repressed inflation in Czechoslovakia, Poland, and Romania between 1980 and 1990.

First, they develop a simultaneous equation model, which stipulates that the repressed inflation is caused by the difference in growth rate between households' money holdings and retail sales in the economy. Following are the model's basic assumptions:

- A stable demand for money function exists where the demand for real quasi-money depends on the level of real income and the expected rate of inflation.
- Inflationary expectations are formed under an adaptive scheme in which expected inflation for the next period depends on the error made in predicting the current period's "true" inflation.
- The real stock of quasi-money balances adjusts to the desired level, after a time lag.

Feltenstein and Ha then derive a reduced-form equation on real quasi-money holdings, and

estimate it with quarterly data for Czechoslovakia and Poland, and with annual data for Romania. Based on the estimated equation, they derive the true inflation rate for each economy. Finally, they determine the significance of the repressed inflation in each economy through statistical tests on parameters of the estimated equation.

These are their main findings:

- During 1980-90 in Czechoslovakia, repressed inflation was insignificant, the demand for money was mostly for transaction purposes, and inflationary expectations were almost myopic.
- In Poland, repressed inflation was significant but decreasing after 1987, and inflationary expectations adjusted fairly rapidly.
- In Romania, repressed inflation was serious throughout the period, and inflationary expectations adjusted quite rapidly.

This economic model needs to be refined and the data used need to be improved. But the paper's findings are broadly consistent with the results of most other studies.

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# **AN ANALYSIS OF REPRESSED INFLATION IN THREE TRANSITION ECONOMIES\***

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# An Analysis of Repressed Inflation in Three Transition Economies<sup>1</sup>

## I. Introduction

A major issue in many centrally planned economies is how to coordinate growth in the money supply with a controlled price system. In addition, the foreign sector is not open and the exchange rate, like the domestic price system, is subject to distortions. As a result, changes in the money supply which are not accompanied by changes in domestic prices and the exchange rate may lead to repressed market demands and balance of payments disequilibrium.

In this paper we will look at the experiences of three countries, Czechoslovakia, Poland, and Romania. The common experience of these countries is, of course, the abandonment of central planning. We see that each of these countries has, to at least some extent, inflationary difficulties. The aim of this paper is to estimate a measure of the extent to which the price level has been repressed because of the years of price controls. We develop a simple analytical model which derives a "true" rate of inflation on the basis of the different rates of change of the stock of money in circulation and the nominal value of retail sales. This true rate of inflation is then used to explain changes in quasi-money balances. We estimate this model for both Czechoslovakia and Poland using quarterly data. We use annual data for Romania, in the absence of quarterly numbers. As we shall see, the estimates for these countries are strikingly different.

Is it possible that prices, which are still subject to many controls, have lagged behind money growth sufficiently so as to create excess nominal demand for goods? Is consumer demand for quasi-money balances merely a function of excess monetary expansion, which forces savings, or is it an outcome of anticipated

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<sup>1</sup>This paper was funded by a World Bank project on formerly planned economies. We would like to thank Jong-goo Park for suggesting the topic, and Fabrizio Coricelli and Adnan Mazarei for helpful discussions.

inflation as well as real income? Finally, is it possible to use observed variables to derive a measure of the unobserved "true" rate of inflation, if prices are indeed, repressed? This question has been addressed in Feltenstein and Farhadian (1987) and Portes and Santorum (1987) as well as others. The approach we use here is based upon Feltenstein, Lebow, and Van Wijnbergen (1990), and, in particular, Feltenstein and Ha (1991).

Let us now turn to a simple model that may help to answer the questions we have posed. The next section will derive our analytical model, while Section 3 will present estimation results. Section 4 will derive certain conclusions concerning the extent to which the official price index currently reflects repressed markets.

## II. The Model

We will suppose that the representative consumer's demand for speculative and precautionary money balances, given by the stock of quasi-money, is a function of his real income and expectations of the rate of inflation. The rate of inflation he expects may not, however, correspond to the expected change in the official price index, since these controlled prices do not necessarily reflect the shortages that he may face. We should note that this model does not incorporate the possibility of the foreign sector as an outlet for excess demand. For the period of our sample, this is a reasonable omission, given the restrictions on foreign trade imposed by the countries in question. As the countries open their foreign sector, our model would become less relevant.

Let  $\pi_T$  denote the "true" rate of inflation as perceived by the consumer. That is, the demand for real quasi-money balances,  $qm^d$ , is then given by:

$$\log qm^d = a_0 + a_1 \log y + a_2 \pi \frac{E}{T} \quad (2.1)$$

where:

$$qm = QM/P_T \quad (2.2)$$

Here QM denotes quasi-money balances, y represents real income

and  $\pi_T^E$  represents the expected rate of inflation in the true price index  $P_T$ , where the true rate of inflation is given by:

$$\pi_T = \log P_T - \log P_{T-1} \quad (2.3)$$

Thus, the true rate of inflation in equation (2.1) may be interpreted as that rate that would induce consumers to hold the observed level of quasi-money balances in the absence of price controls. Our aim will be to determine whether or not this is significantly different from the official inflation rate.

Why do we deflate quasi-money balances by the true price index rather than the official price index since, in the absence of black-markets, the official index represents the cost of current consumption? Suppose that the consumer expects the reform to succeed and hence that markets will fully clear in subsequent periods. In this case  $\pi_T$  will reflect the cost of future consumption, since it reflects the consumer's best estimate of inflation in the absence of price controls. If, on the other hand, the consumer believes that price controls will never be lifted, then he should use the official price index,  $\pi$ , as a measure of the cost of future consumption. In deflating by the true price index we are therefore assuming that the consumer expects price controls to be lifted. Although we have not done so here, one can test the alternative hypotheses of future price decontrol versus no decontrol.<sup>2</sup>

The normal specification of equation (2.1) would be to have the dependent variable be the total real money supply, that is  $(QM+M_1)/P$ . Here we are using an analogue of a Baumol-Tobin specification in which the consumer chooses to divide his cash holdings between bank deposits and currency. If there is increasing excess demand, and hence in our model increased repressed inflation, then the "true" real interest rate on these savings deposits declines. As the "true" rate of inflation rises, then the relative gap between the real returns

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<sup>2</sup>Feltenstein, Lebow, and Van Wijnbergen (1990) carry out such a test for China and determine that the hypothesis of future decontrol is preferred to that of permanent controls.

on currency and on quasi-money deposits narrows, even though the real return on quasi-money always remains higher. The consumer can hold his assets in the form of currency, quasi-money, or goods, but the presence of repressed demand makes increased holdings of goods impossible to achieve immediately. Since goods become available sporadically, and are instantly purchased, it is essential that the consumer have cash readily available to make purchases. Interest bearing deposits are not fully liquid, and so will not be available if the consumer needs to suddenly make a purchase. Currency thus becomes relatively more attractive to the consumer than quasi-money deposits, leading him to reduce his interest bearing deposits.<sup>3</sup> Thus the liquidity differences between currency and quasi-money deposits cause the consumer to be willing to forego the slightly higher real returns offered by quasi-money.

We use the ratio of the broad money balances to retail sales to measure the deviation between "true" and official inflation. Here broad money is defined as the sum of currency and consumer holdings of bank deposits. The idea is that an increase in the ratio of the rates of growth of broad money and retail sales may reflect the existence of repressed demands caused by price controls.<sup>4</sup> Let  $M_2$  be the broad money balance, which is equal to stock of currency in circulation plus households bank deposits,  $P$  be the official price level and  $R$  be the real volume of consumer retail sales. We then set:

$$\log P_T = \log P + \alpha \log (M_2/PR) \quad (2.4)$$

Thus if  $\alpha = 0$  then the true and official rates of inflation are equal. If, on the other hand,  $\alpha = 1$ , for example, then a 10 percent increase in the ratio of money to retail sales will cause the true rate of inflation to be 10 percent

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<sup>3</sup>It should be noted that the phenomenon of substitution of currency for quasi money occurs in many high inflation countries even without the presence of price controls. Thus, for example, in the 1923 German inflation the real stock of fixed-interest rate deposits fell much more rapidly than did the real stock of  $M_1$ .

<sup>4</sup>This does not mean, however, that an increase in money, in the absence of price controls, must go entirely into consumption.



higher than the official rate. Thus we use a measure of monetary overhang to determine the difference between true and official inflation rates.<sup>5</sup> It should be noted that we would expect that  $\alpha \leq 1$  so the equation (2.4) may be written as:

$$\log p_T = \alpha \log(M/R) + (1 - \alpha) \log p$$

or:

$$\pi_T = \alpha \Delta \log(M/R) + (1 - \alpha) \pi$$

Thus if  $\alpha \leq 1$  any increase in official prices, that is, an increase in the current price level or, equivalently, an increase in the current rate of inflation, would cause an increase in the rate of inflation in the true price index. However unless  $\alpha = 0$ , that is, no repressed inflation, the increase in  $\pi$  would cause a smaller increase in  $\pi_T$  (as  $\alpha \leq 1$ ), thus reducing the divergence between the two price indices and hence reducing the degree of repression in the price level. If  $\alpha > 1$  then the above specification would be difficult to interpret. Our estimations indicate, however, that  $\alpha$  is less than 1.

Implicitly, a value of  $\alpha > 0$  indicates that monetary overhang leads to aggregate excess demand. Walras' law ensures that the value of excess demand is equal to zero, so there must be excess supply in some markets, as well as excess demand in others. It is therefore plausible that there might be price deflation in those markets with excess supply. If, on the other hand, prices are rigid downwards, then we would not observe sectoral deflation. In any case, given the aggregate nature of our model we are unable to investigate sectoral price changes.

We will assume that inflationary expectations are determined according to an adaptive pattern in which the expected value of the true rate of inflation for next period depends upon the error made in predicting the true rate of inflation for this period. Thus:

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<sup>5</sup>Feltenstein and Farhadian (1987) use a simpler measure in which there is a constant factor between the true and official inflation rates. Another approach to the specification of repressed inflation in China is given in Portes and Santorum (1987).

$$\pi_T^E - \pi_{T-1}^E = \beta(\pi_T - \pi_{T-1}^E) \quad (2.5)$$

Here  $\beta$  is an estimated constant that we hypothesize lies between 0 and 1. In addition, we will suppose that the real stock of quasi-money balances, valued at the true price level, adjusts with a lag to the desired level. Accordingly we have:

$$\log qm - \log qm_{-1} = \tau (\log qm^d - \log qm_{-1}) \quad (2.6)$$

where  $\tau$  is a parameter to be estimated.

Thus the economic interpretation of our model is that the stock of currency plus household bank deposits is exogenous and represents nominal purchasing power. Consumers then try to adjust their interest-bearing deposits on the basis of their anticipations of the true rate of inflation, as well as income. It seems plausible that the selection of quasi-money balances would be a consumer choice variable, although aggregate money issue would be determined by the government.<sup>6</sup>

### III. Estimations

We will consider the three countries of our study in order.

#### a. Czechoslovakia

We use quarterly data from 1980 to 1990. To estimate the model, we need 5 series: money, quasi-money, consumer price index, household income, and households purchase of consumption goods. Our data sources are as follows (all data is taken from International Financial Statistics):

- a) Money (M1) is defined as currency outside banks plus demand deposits.
- b) Quasi-money: Quarterly data is available only from 1988.1 to 1990.4.

However annual data is available from 1980. We assume that the seasonal pattern

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<sup>6</sup>See De Wulf and Goldsborough (1986) as well as Feltenstein and Farhadian (1987) for descriptions of the process of monetary creation in China.

in each year prior to 1988 was the same as the average seasonal pattern of 1988-1990. We compute the average seasonal pattern of 1988-1990, and use it together with the annual data to generate the quarterly data for the period from 1980 to 1987. In this case, as in the cases of the two other countries of our study, we have a problem with data coverage. Available data for quasi-money balances does not distinguish between personal and corporate deposits. Therefore our data, which should reflect only personal deposits, may be contaminated by corporate deposits. Anecdotal evidence indicates that corporate bank deposits are largely in the form of checking deposits, and hence would not effect our results. In any case, until greater detail in the monetary accounts becomes available, we have no alternative to using the current definition of quasi-money.<sup>7</sup>

c) Consumer price index (CPI): This series is directly available on a quarterly basis.

d) Household income (Y): Neither annual nor quarterly data is available for this variable. We therefore have to use a proxy as the scale variable for the estimation of money demand. We choose GNP for the following two reasons. First, it is reasonable to believe that GNP and households income are highly correlated. Second, we do not have quarterly data of for any acceptable scale variable other than GNP.

Quarterly data from 1988 to 1990 is available, while annual data is available for prior years. The quarterly data prior to 1988 is then constructed in the same way as in b).

e) Households purchase of consumption goods (R): Unfortunately, we do not have any quarterly data for this variable, although annual data is available. We therefore assume that the marginal propensity to consume is constant, so that household consumption follows the same seasonal pattern as households income. Based on this assumption, the quarterly data is constructed in the same way as in b).

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<sup>7</sup>In previous work on China, Feltenstein and Ha (1991), we determined that corporate deposits were a stable fraction of broad money over a long time period. One might expect the same phenomenon to occur in the case of these three countries, although we do not have direct evidence to that effect.

From the complete model given by equations (2.1) - (2.6) we derive a reduced form for estimation:

$$\begin{aligned}
 q_{mm} &= A_0 + A_1 Y Y + A_2 \pi_T + A_3 q_{mm}_{-1} \\
 \text{where } q_{mm} &= \log q_m - (1 - \beta) \log q_{m-1} \\
 Y Y &= \log y - (1 - \beta) \log y_{-1} \\
 \pi_T &= \pi + \alpha (\log (M_2/PR) - \log (M_2/PR)_{-1}) \\
 \log q_m &= \log QM - \log P - \alpha \log (M_2/PR) \\
 A_0 &= \tau a_0 \beta \quad A_1 = \tau a_1 \quad A_2 = \tau a_2 \beta \quad A_3 = 1 - \tau
 \end{aligned}$$

We search for the values of  $\alpha$  and  $\beta$  to maximize the likelihood function, subject to the constraint:  $1 < \beta$ ,  $\tau < 1$ . We find that  $\alpha = 0.09$  and  $\beta = 0.08$ . The estimation of the reduced form is:

$$\begin{aligned}
 q_{mm} &= 0.323 + 0.570yy - 0.108\pi_T + 0.367q_{mm}_{-1} \\
 &\quad (5.717) \quad (2.542) \quad (-0.476) \quad (3.287) \\
 R^2 &= 0.631, \quad H = 2.610, \quad \log\text{-likelihood} = 123.248
 \end{aligned}$$

The numbers in parentheses denote t-statistics. All coefficients have the correct sign, and, except for the inflation rate, are significant at the 1 percent level. The parameters of the structural form can be solved as

$$a_0 = 6.378, \quad a_1 = 0.900, \quad a_2 = -2.132, \quad \tau = 0.633.$$

The estimate  $\tau$  indicates that real quasi-money balances take about  $(1-\tau)/\tau = 0.58$  quarters to adjust to the difference between the demand for quasi-money and the actual stock in the previous quarter.

The estimate  $\alpha = 0.09$  indicates that when the growth rate of money stock is 10 percent higher than that of the nominal retail sales, the true rate of inflation will rise by only 0.9 percent above the reported rate of inflation.

Intuitively, this implies that there was no significant repressed inflation during the sample period.

We may carry out a formal statistical test of repressed inflation hypothesis in the following way. The null hypothesis is that there is no repressed inflation, i.e.,  $H_0: \alpha = 0$ . The value of the log-likelihood function under  $H_0$  is 122.5699. The chi-square statistic is 1.356, while the 0.9 value of the chi-square is 2.71. Therefore, we can not reject the null hypothesis.

The parameter  $\beta = 0.08$  indicates that the expectation of inflation appears to be revised fairly slowly. A formal statistical test fails to reject the hypothesis of myopic expectations ( $H_0: \beta = 0$ ).

Another interesting finding is that the coefficient of the true inflation rate,  $\pi_T$ , is statistically insignificant. In fact the velocity of money is almost constant, so that the official and the true inflation rates are virtually identical (see Figures 1 and 2). The constant velocity thus indicates that inflation is not used to measure the opportunity cost of money holding.<sup>8</sup> At the same time, deposit interest rates were fixed at 3 percent from 1970 to 1990. Our results thus lead to the conclusion that the demand for money in Czechoslovakia is purely for transaction purposes.

What is the intuition behind our conclusion of a fixed true velocity of money in Czechoslovakia? In a typical developed country we expect to see significant fluctuations in the velocity. In the United States, for example, we see a rising velocity. Czechoslovakia, like most planned economies, has lacked the variety of financial instruments that permit substitution out of money. Accordingly, the conclusion of a constant velocity is not unexpected. In particular, the absence of alternative assets means that standard measures of the opportunity cost of holding money would not be relevant.

Our conclusions are summarized as follows:

(1) There was no repressed inflation or forced savings in the Czechoslovakia during the period from 1980 to 1990.

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<sup>8</sup>We should note that our model does not assume that the officially measured velocity of money is constant. In this case, however, both the officially measured velocity, as well as our "true" velocity, are approximately constant.

(2) Demand for money was only for transaction purposes.

(3) Inflationary expectations were almost myopic.

These results support the general belief that Czechoslovakia has had a long history of conservative economic management. Indeed, most current estimates would claim that total inflation during the 1980's, including any repressed inflation, averaged only about 5 percent per year.<sup>9</sup> It is useful to examine price developments since 1990, the end of our sample period, to see if subsequent events tend to support our conclusion that there was little repressed inflation in Czechoslovakia.

The economy was liberalized in a single "cold-turkey" program on January 1, 1991, in which prices covering about 85 percent of the value of total transactions were liberalized. At the time, it was anticipated that inflation would rise to about 30 percent. This would be caused largely by the shift to world prices among CMEA members. Tight fiscal and monetary policies were to simultaneously be imposed. In fact, prices rose by about 50 percent at the beginning of 1991, rather than the targeted 30 percent. The main cause of this outcome seems not to have been repressed inflation, however, but a drastic fall in GDP which reduced tax revenues and created inflationary pressures via budget imbalances. After credit policy was adjusted to compensate for the unexpected price increases, inflation was virtually eliminated for the second half of 1991.

It may be useful, in this context, to note a recent paper, Drabek, Janacek, and Tuma (1992), that surveys current literature on inflation in Czechoslovakia. Examining the period 1985-91, the general conclusion of a number of the relevant articles is that hidden, or repressed, inflation ranged from about 0.5 to 2 percent annually in the consumer market. These figures are so low as to give support to our conclusion that there was no statistically significant repressed inflation in Czechoslovakia.

We may thus conclude that casual evidence for 1991 supports our contention that repressed inflation has not been a problem for Czechoslovakia. We have also

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<sup>9</sup>See, for example, The Czech and Slovak Federal Republic, an Economy in Transition (1990).

considered the question of whether or not the adaptive expectations framework might be improved upon, that is, would a different estimation approach lead to different conclusions. Accordingly, we have carried out an alternative estimation, using a technique similar to that developed in Lane (1992) for Poland. The results suggest also that there was no significant repressed inflation in Czechoslovakia during the sample period and the expectations of inflation were myopic.<sup>10</sup>

Let us now turn to the quite different case of Poland.

b. Poland<sup>11</sup>

As in the case of Czechoslovakia, we estimate the model for the sample period from the first quarter of 1980 to the second quarter of 1990.

To estimate the model we need five time series:

- consumer price index (CPI),
- money ( M1 = currency plus demand deposits),
- quasi-money (QM),
- retail sales (RETSALE),
- household income (Y).

a) CPI: This series (from 1980.1 to 1990.2) is directly available.

b) M1: the data from 1983.1 to 1990.2 is directly obtained from International Financial Statistics, a TSP file provided by the World Bank, and Historically Planned Economies: A Guide to the Data (1992). Data for the fourth quarters of 1980 to 1982 is also available from IFS. We then generate quarterly data for 1980 to 1982 by using the seasonal patterns of 1983-90.

c) QM: We have quarterly data from 1986.1 to 1990.2, and the data of the fourth quarters of 1980 to 1985. Using the seasonal pattern of 1986, we construct the quarterly data from 1980 to 1985.

d) RETSALE: the index of nominal retail sales (from 1983.1 to 1990.2)

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<sup>10</sup> See Annex 1 for details.

<sup>11</sup>We would like to thank Fabrizio Coricelli for providing us with the data used for our estimations.

is given by data supplied by the World Bank. The growth rates over 4 quarter periods from 1983 to 1987 are almost constant, and then increase rapidly after 1987. We have therefore chosen to assume that the growth rate of retail sales from 1980 to 1982 was similar to that from 1983 to 1987. Therefore, we construct the seasonal data from 1980 to 1982 by assuming that the growth rate was the same as the average growth rate from 1983 to 1987.

e) Y: This series is not available. To find a proxy to this series, we follow the same procedure as in Feltenstein and Ha (1991), where household income is assumed to equal retail sales (or private consumption) plus changes in household asset holding. This assumption is reasonable in centrally planned economies because other financial assets, such as stocks or bonds, are not available and the only means of saving is to hold currency and bank deposits.

To employ this procedure, we must have data on nominal retail sales. Unfortunately only its index number is available. However, retail sales to the private sector constitute the main part of private consumption. The data we need is the quantity of money transactions. Since personal credit and private loans were not available in centrally planned economies, retail sales to the private sector are essentially equal to private consumption. The only difference between these two variables would normally be that private consumption includes farmers' self-produced goods as well as retail purchases. However, the share of self-produced goods in private consumption is very small. Using China, where 75 percent of the total labor force is rural, as an example, retail sales to the private sector accounts for 95 percent of private consumption. We therefore believe that using private consumption as a proxy for retail sales will not change our results.

The annual data for private consumption is available. We thus construct quarterly data for private consumption, using the seasonal pattern of retail sales.

We search for the values of  $\alpha$  and  $\beta$  to maximize the likelihood function, subject to the constraint:  $1 < \beta$ ,  $\tau < 1$ . We find that  $\alpha = 0.58$  and  $\beta = 0.84$ .



The estimation of the reduced form is:

$$\begin{aligned} qmm &= -0.359 + 0.363yy - 0.323\pi_T + 0.744qmm_{-1} \\ & \quad (-4.109) \quad (5.575) \quad (-8.818) \quad (8.745) \\ R^2 &= 0.956, \quad H = 0.570, \quad \log\text{-likelihood} = 71.206 \end{aligned}$$

The numbers in parentheses denote t-statistics. All coefficients have the correct sign, and are significant at the 1 percent level. The parameters of the structural form can be solved as

$$a_0 = -1.402, \quad a_1 = 1.418, \quad a_2 = -1.262, \quad \tau = 0.256.$$

The estimate  $\tau$  indicates that real quasi-money balances take about  $(1-\tau)/\tau = 2.9$  quarters to adjust to the difference between the demand for quasi-money and the actual stock in the previous quarter.

The estimate  $\alpha = 0.58$  indicates that when the growth rate of money stock is 10 percent higher than that of the nominal retail sales, the true rate of inflation will rise by 5.8 percent above the reported rate of inflation. This implies that there was repressed inflation during the sample period. Figure 3 shows the deviation between the true and the official price index during the period.

Formal statistical test of repressed inflation hypothesis is carried out as follows. The null hypothesis is that there is no repressed inflation, i.e.,  $H_0: \alpha = 0$ . We therefore set  $\alpha = 0$  and search for a value for  $\beta$  by using the maximum likelihood technique mentioned above. The value of log-likelihood function under  $H_0$  is 49.045. The chi-square statistic is 44.322, while the 0.99 value of chi-square is 9.21. Therefore, the null hypothesis is decisively rejected.

The parameter  $\beta = 0.84$  indicates that the expectation of inflation appears to be revised fairly rapidly, as might be expected in an economy with a rapidly changing institutional structure as well as a rising inflation rate. Formal

statistical test rejects the hypothesis of myopic expectations ( $H_0: \beta = 0$ ), and the hypothesis of perfect foresight ( $H_0: \beta = 1$ ). This confirms the specification of adaptive expectations.

The official velocity of money,  $V = PR/M$ , where  $P$  is the officially reported price series, was rather smooth prior to 1987, but it has been fluctuating since 1987. We define the true velocity of money as  $V_T = P_T R/M$ , where  $P_T$  is the true price level. As pointed out by Feltenstein and Ha (1991), the hypothesis of constant true velocity is equivalent to  $\alpha = 1$  in this model. Setting  $\alpha = 1$ , the value of log-likelihood function equals 60.821. The hypothesis of constant velocity is then statistically rejected. This conclusion is thus an example of how our approach does not require a constant "true" velocity of money. Rather, it presents a testable hypothesis which, in this case, is rejected. Finally, Figure 4 compares true and official inflation rates

This result is thus quite different from Feltenstein and Ha (1991) who applied a similar model to China. They found that the officially measured velocity was decreasing and the true velocity was constant (i.e.,  $\alpha = 1$ ) in China during 1979 to 1988. Since the estimated value of  $\alpha$  is smaller in Poland than in China, one may conclude that the extent of repressed inflation is weaker in Poland than in China. Alternatively, this finding may reflect the fact that greater price liberalization has occurred in Poland than in China, or, as indicated in Lane (1992), that there is considerable substitution between money and black-market assets in Poland, while there may be less such substitution in China. Indeed, the presence of alternative black-market assets may explain the difference from our results on Czechoslovakia.

Another interesting finding is that the coefficient of inflation rate is statistically significant, even if the official rate of inflation is used. That is, the coefficient of  $\pi_T$  is significant even if  $\alpha$  is set to 0. This result indicates that the official rate of inflation plays some role in household demand for real balances, although it is not as satisfactory a determinant of money demand as the true rate of inflation. This is different from the results for China, where official inflation rate does not play a role, but true inflation rate does.

It is also different from the results for Czechoslovakia where neither official nor true inflation rates play a role in determining money demand.

Our conclusions are summarized as follows:

(1) There was repressed inflation and forced savings in Poland during the period from 1980 to 1990. However, the officially reported price level has been converging to the true price level since 1987.

(2) Inflationary expectations were neither myopic nor did they exhibit perfect foresight, although expectations adjusted fairly rapidly.

(3) Neither the official nor the true velocity of money was constant. Indeed, velocity has been fluctuating significantly since 1987.

(4) The true rate of inflation, estimated by the model, performs better significantly better than the official rate as a measure of the opportunity cost of money holding. The official rate also plays a role in determining money demand.

It may be useful to compare our findings with those of a recent paper by Lane (1992). Here an error correction estimation technique is used to estimate a money demand model for Poland. Unlike the narrow range of explanatory variables used in our paper, Lane's approach incorporates information on black markets in commodities and in foreign exchange. Thus it is difficult to make direct comparisons, although certain similarities do emerge. Lane concludes that household money balances adjust fairly rapidly toward their equilibrium level. Our conclusion, reflected in the estimated value of  $\tau$  is that quasi-money balance adjust to their equilibrium level as consumers substitute between money and quasi-money. Thus, as in Lane, money holdings are not passive, although we cannot replicate his findings about substitution between money and black-market assets.

Let us briefly consider if recent events in Poland support our conclusions. Beginning in 1989 the Polish government began to reduce direct controls over a gradually increasing spectrum of prices, and by the end of 1990 almost 85 percent of retail prices were freely determined. At the same time there was, initially, an extraordinary burst of inflation. In the fourth quarter of 1989 the wholesale

price index rose by 65 percent over the previous quarter, while in the first quarter of 1990 it rose by an additional 550 percent. Prices then stabilized and in the second quarter of 1990 the wholesale price index increased by only 1.5 percent. This trend continued in the third quarter of the year, while in the fourth quarter a relaxation in monetary policy caused inflation to increase to a monthly rate of about 5 percent.

We may thus conclude that there appears to be evidence supporting our notion of repressed inflation. Indeed, it would seem that price liberalization brought about a once-and-for-all price adjustment, at the beginning of 1990, that eliminated monetary overhang and corresponding repressed inflation.

Let us now turn to Romania, the final country in our study.

c. Romania

Unlike Czechoslovakia and Poland, quarterly data is not available, at least to us, for Romania. We have therefore estimated our model using annual data from 1975 to 1990.

In order to estimate the model we need the same five series as before:

consumer price index (CPI),  
money ( M1 = currency plus demand deposits),  
quasi-money (QM),  
retail sales (RETSALE),  
household income (Y).

a) CPI, M1 and QM: The time series for CPI, M1 and QM are directly obtained from World Bank data.

b) RETSALE: the series retail sales is given in an unpublished World Bank document.

c) Y: This series is not available. To find a proxy to this series, we follow the same procedure as in Feltenstein and Ha (1991), where household income is assumed to equal retail sales (or private consumption) plus the change in household asset holding. This assumption is justified in the same manner as in the case of the previous countries.

We search for the values of  $\alpha$  and  $\beta$  to maximize the likelihood function, subject to the constraint:  $1 < \beta$ ,  $\tau < 1$ . We find that  $\alpha = 1.14$  and  $\zeta = 0.91$ . The estimation of the reduced form is:

$$\begin{aligned}
 qmm &= -0.651 + 0.921yy - 1.333\pi_T + 0.280qmm_{-1} \\
 &(-7.199) \quad (8.632) \quad (-8.885) \quad (3.508) \\
 R^2 &= 0.964, \quad H = 0.107, \quad \log\text{-likelihood} = 35.497
 \end{aligned}$$

The numbers in parentheses denote t-statistics. All coefficients have the correct sign, and are significant at the 1 percent level. The goodness-of-fit and H-statistic are satisfactory. The parameters of the structural form can be solved as

$$a_0 = -0.904, \quad a_1 = 1.279, \quad a_2 = -1.851, \quad \tau = 0.720.$$

The estimate  $\tau$  indicates that real quasi-money balances take about  $(1-\tau)/\tau = 0.4$  quarters to adjust to the difference between the demand for quasi-money and the actual stock in the previous quarter.

The estimate  $\alpha = 1.14$  indicates that when the growth rate of money stock is 10 percent higher than that of the nominal retail sales, the true rate of inflation will rise by 11.4 percent above the reported rate of inflation. This implies that there was high repressed inflation during the sample period, as shown in Figure 5.

Formal statistical test of repressed inflation hypothesis is carried out as follows. The null hypothesis is that there is no repressed inflation, i.e.,  $H_0: \alpha = 0$ . We therefore set  $\alpha = 0$  and search for a value for  $\beta$  by using the maximum likelihood technique mentioned above. The value of log-likelihood function under  $H_0$  is 26.935. The chi-square statistic is 17.124, while the 0.99 value of chi-square is 9.21. Therefore, the null hypothesis is decisively rejected.

The parameter  $\beta = 0.91$  indicates that the expectation of inflation appears to be revised fairly fast. Formal statistical test rejects the hypothesis of myopic expectations ( $H_0: \beta = 0$ ) at 1 percent significance level, but the hypothesis of perfect foresight ( $H_0: \beta = 1$ ) can not be rejected.

The official velocity of money,  $V = PR/M$ , where  $P$  is the officially reported price series, decreased by 70 percent from 1975 to 1990. The following test will show that the decline of velocity is due to the use of official price indices. We define the true velocity of money as  $V_T = P_T R/M$ , where  $P_T$  is the true price level. As pointed out by Feltenstein and Ha (1991), the hypothesis of constant true velocity is equivalent to  $\alpha = 1$  in this model. Setting  $\alpha = 1$ , the value of log-likelihood function equals 35.180. The hypothesis of constant velocity can therefore not be rejected. Figure 6 indicates the extent to which there is a divergence between the true and official velocities of money.

Our conclusions are summarized as follows:

(1) There was severe repressed inflation and forced savings in Romania during the period from 1975 to 1990. The situation does not seem to improve toward the end of the sample period.

(2) The adjustment of inflationary expectations was quite rapid. (3) The true velocity of money was constant, although the official velocity was declining.

Given the limited availability of data for Romania, in particular the absence of quarterly price indices, it is difficult to use the post-sample period to draw inferences about the nature of repressed inflation prior to the beginning of liberalization. It may be useful, however, to quote a recent study by Demekas and Khan (1991) to support some of our conclusions.

Although it is extremely difficult to estimate even roughly the monetary overhang resulting from the fundamental changes in the demand for money taking place in a situation of systemic upheaval and great uncertainty like the one that occurred in Romania in 1990, the pattern of income velocity of money can be indicative. Income velocity of broad money at the end of 1990 in Romania was about half its level during the late 1970's and early 1980's, suggesting that perhaps up to 50 percent of the money stock was involuntarily held.

#### IV. Summary and Conclusion

We have constructed a simple model that tests for repressed inflation by estimating a "true" rate of inflation that explains behavior of observed money demand. We have estimated the model using quarterly data for Czechoslovakia and Poland, and using annual data for Romania. Although our results should be viewed as preliminary, given the imperfect nature of our data, we do have strong evidence that, prior to 1991, there was considerable repressed inflation in Poland and Romania, while there was essentially no repressed inflation in Czechoslovakia.

We intend to further explore this topic by employing alternative methods of estimation. In addition, as more recent data becomes available, we would like to see if there is still any repressed inflation in either Poland or Romania, or if recent high rates of inflation have caused it to vanish, as indeed appears to be the case in Poland.

Figure 1. Czechoslovakia: Official Price (P) and "True" Price (PT)

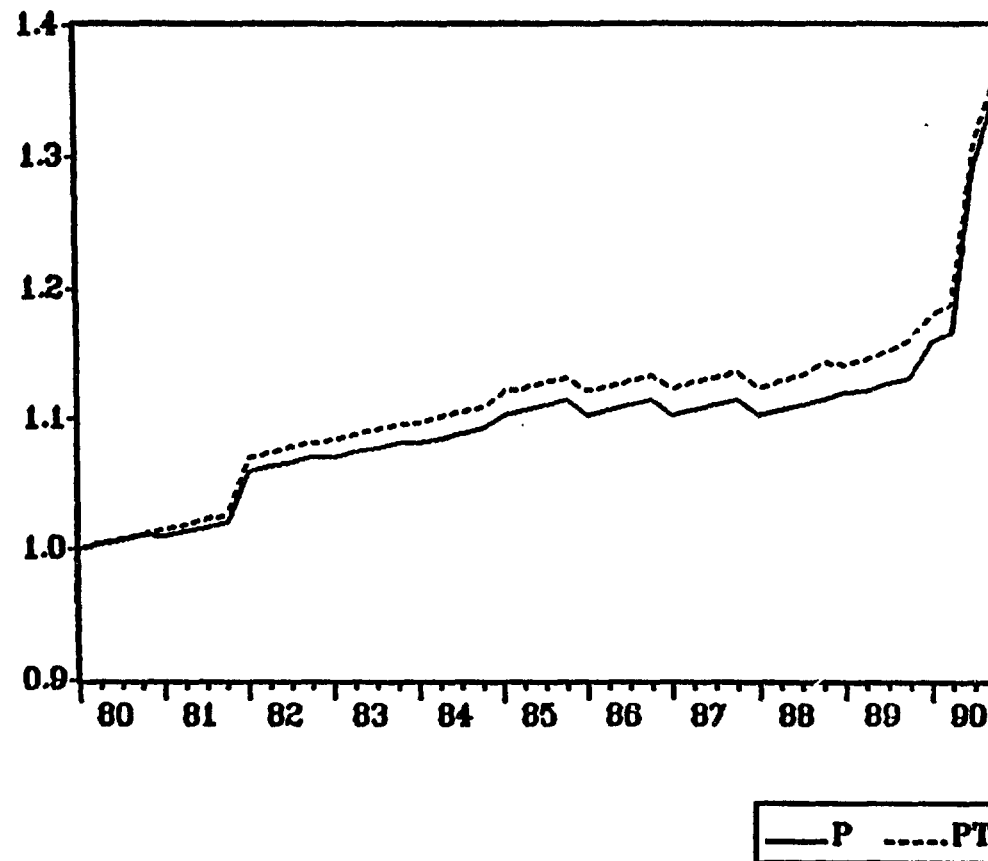




Figure 2. Czechoslovakia: Official and 'True' velocities (V and VT)

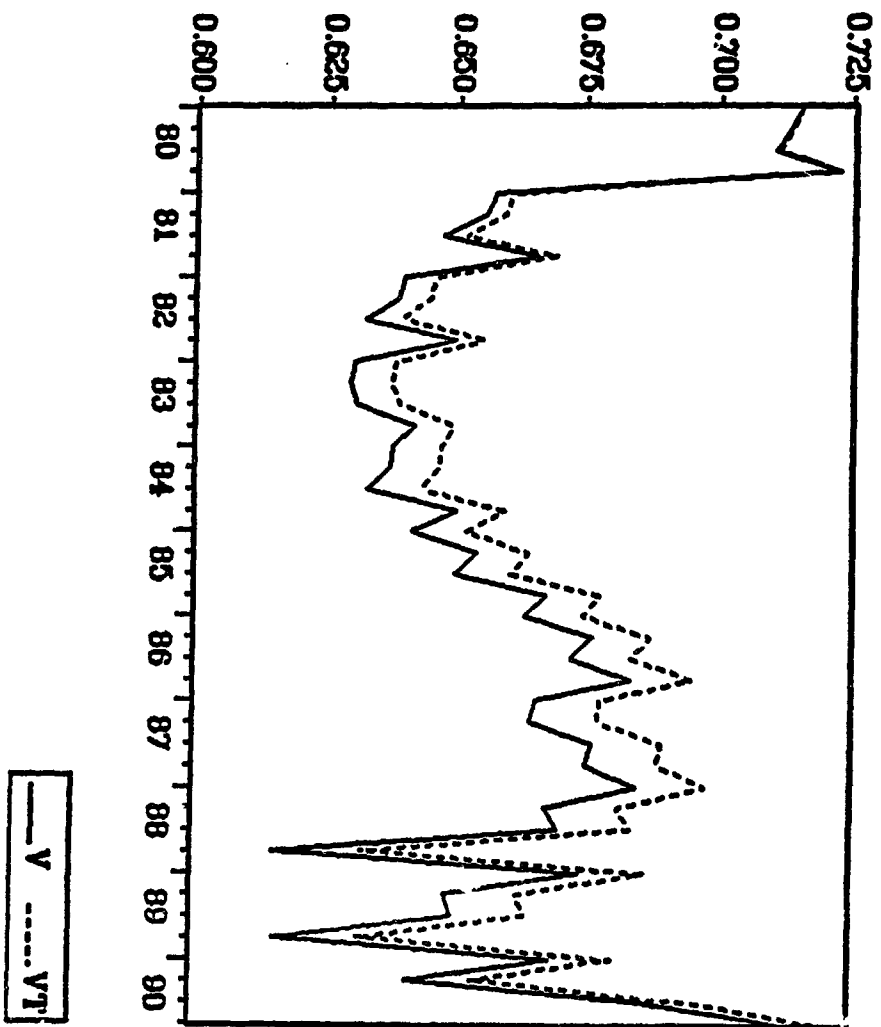


Figure 3. Poland: Official price (P) and "True" Price (PT)

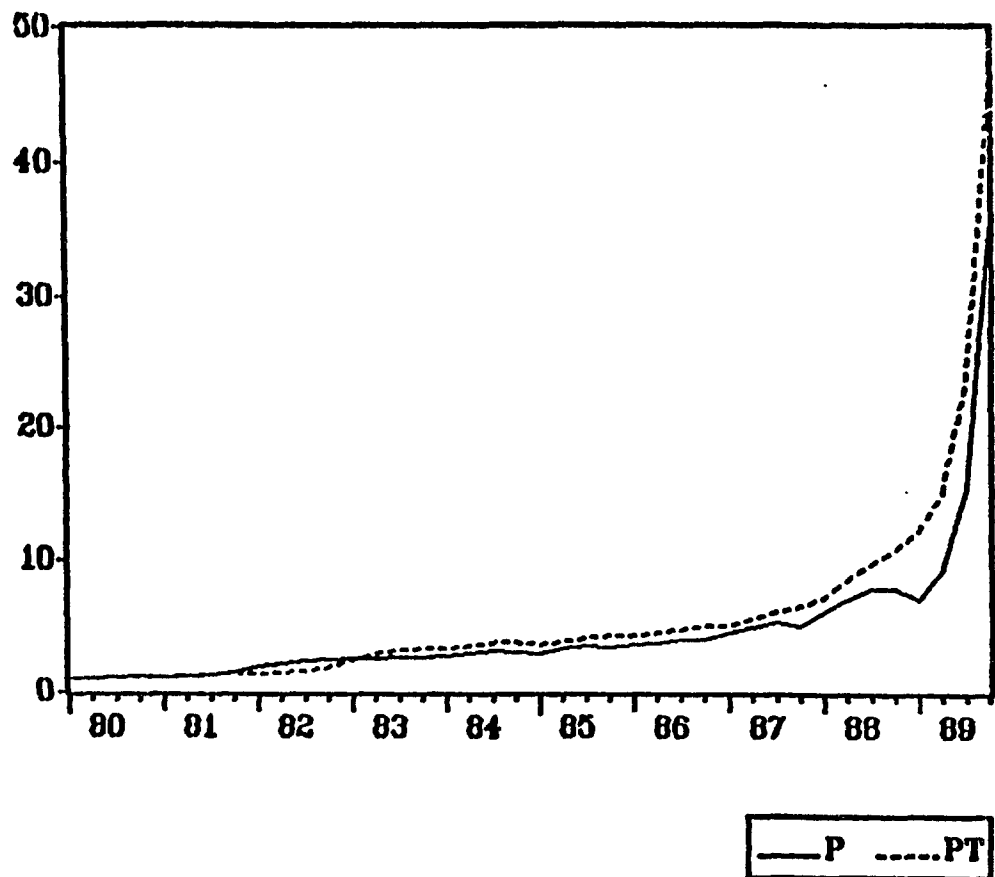


Figure 4. Poland: Official and "True" Velocities (V and VT)

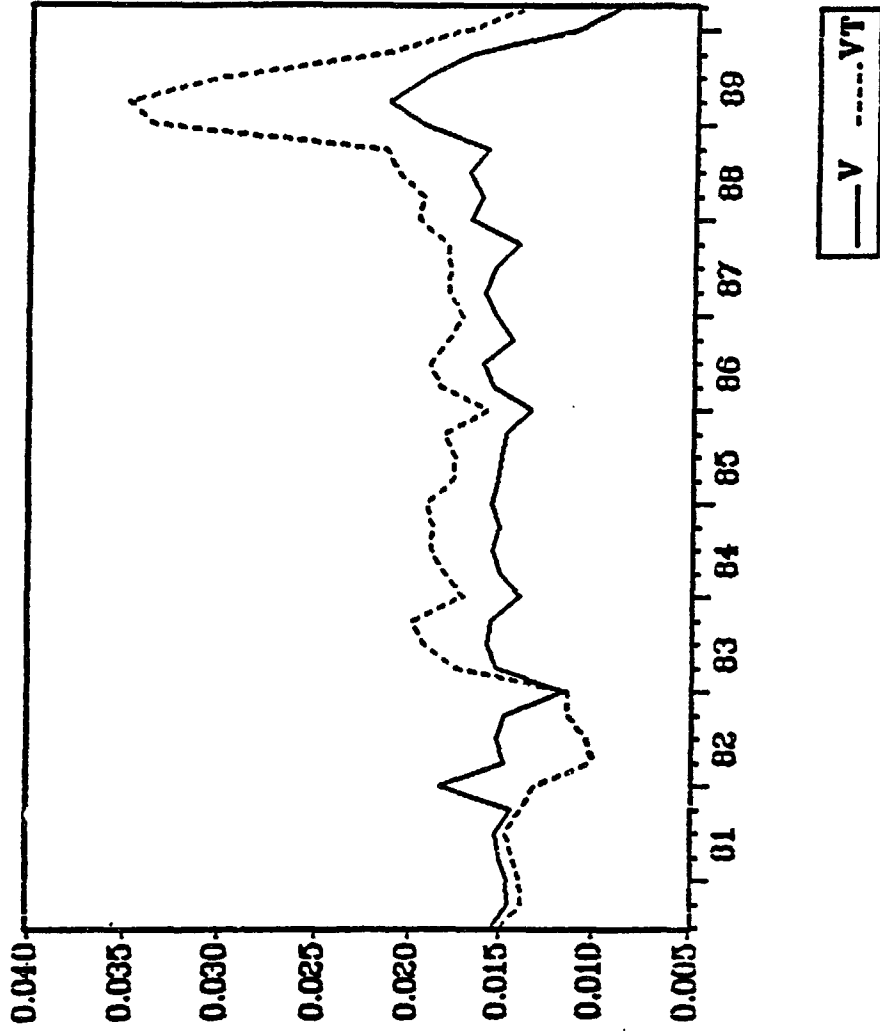


Figure 5. Romania: Official Price (P) and "True" Price (PT)

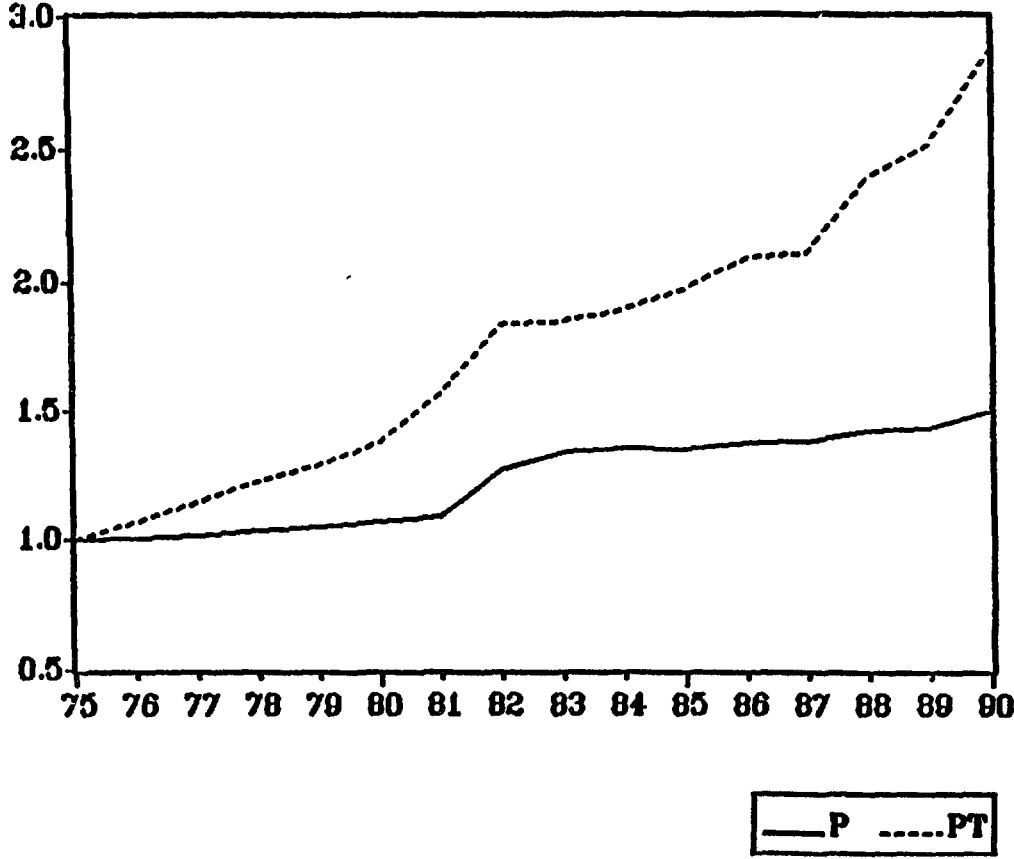
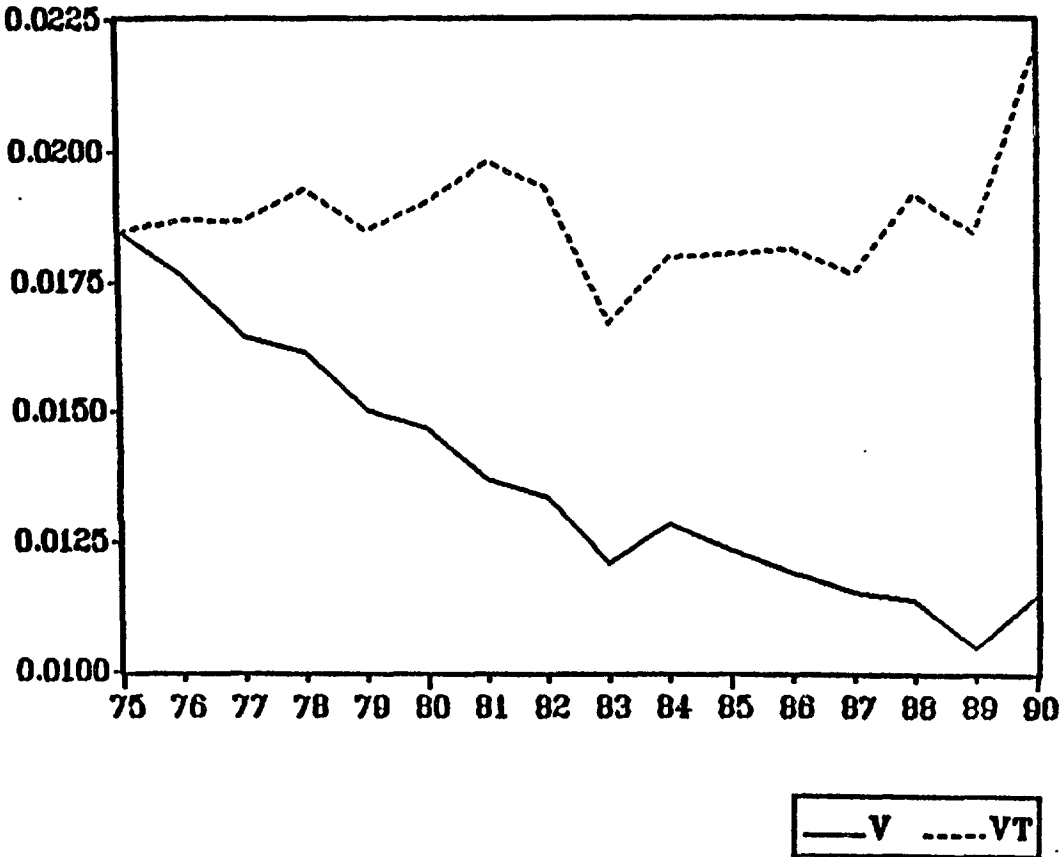


Figure 6. Romania: Official and "True" Velocities (V and VT)



### Estimation of Inflation through Error Correction Model

After the true inflation rate is derived, as in the framework of Feltenstein and Ha (1991), an error correction model (ECM) (see Hendry and Richard (1982)) is estimated. The ECM encompasses the partial adjustment model, and allows more general dynamic adjustment schemes (see Burton and Ha (1990)).

The implementation of this approach first involves testing down from a general autoregressive distributed lag model of the form,

$$M_t = \alpha_0 + \sum_{i=0}^n \beta_i Y_{t-i} + \sum_{i=0}^n \gamma_i Z_{t-i} + \sum_{i=0}^n \delta_i M_{t-i} + v_t \quad (3.1)$$

where  $M$  is the log of real quasi-money balances,  $Y$  is the log of a scale variable, and  $Z$  is an opportunity cost variable, through the sequential deletion of insignificant lags on the basis of an F test. The resulting dynamic model may then be rewritten in error correction form. For example, if the tested-down model is given by

$$M_t = \alpha_0 + \beta_0 Y_t + \beta_1 Y_{t-1} + \gamma_0 Z_t + \gamma_1 Z_{t-1} + \delta_1 M_{t-1} + v_t \quad (3.2)$$

then this may be reformulated as

$$\Delta M_t = \alpha_0 - (1 - \delta_1) \left[ M_{t-1} - \frac{(\beta_0 + \beta_1)}{(1 - \delta_1)} Y_{t-1} - \frac{(\gamma_0 + \gamma_1)}{(1 - \delta_1)} Z_{t-1} \right] + \beta_0 \Delta Y_t + \gamma_0 \Delta Z_t + v_t \quad (3.3)$$

The term in square brackets is the deviation of real balances in period  $t-1$  from their long-run equilibrium value; the coefficient  $(1 - \delta_1)$  captures the responsiveness of real balances to such deviations, and must be positive for

stability. The final step in the process is to estimate equation (3.3), using the deviations from long-run real balances calculated from the tested-down lag model as the error correction variable in square brackets.

We estimate the model for Czechoslovakia, and find the following results

$$\begin{aligned} \Delta q_{mm} &= 0.001 - 0.260 \text{err}_{-1} + 0.270 \Delta y y - 0.581 \Delta \pi_T \\ &(0.923) \quad (-2.268) \quad (2.116) \quad (-4.211) \\ R^2 &= 0.768, \quad D.W. = 2.244, \quad \log\text{-likelihood} = 136.573 \end{aligned}$$

To compare the ECM with the partial adjustment model, we rewrite equation (3.2) as

$$\Delta M_t = (1 - \delta_1) (M_t^* - M_{t-1}) - \beta_1 \Delta Y_t - \gamma_1 \Delta Z_t + v_t \quad (3.4)$$

where  $M^*$  can be calculated from the estimated long-run model. It follows that the partial adjustment model implies the restriction  $\beta_1 = \gamma_1 = 0$ .

We then estimate (3.4), and test the hypothesis of a partial adjustment scheme. The chi-square statistic is 25.72, while the 0.99 value of the chi-square is 9.21. The result is thus in favor of the ECM. Some aspects, however, merit comment. We estimate the ECM using the series of the true rates of inflation that are estimated from the partial adjustment model. A formal test of the ECM against partial adjustment models would involve estimating the true rate of inflation in both models. However, due to the complexity of our theoretical model, it is difficult, if not impossible, to estimate the parameters  $\alpha$  and  $\beta$  in an ECM. On the other hand, our estimates and tests suggest that there was no repressed inflation in Czechoslovakia during the sample period ( $\alpha = 0.09$ ), and the expectations of inflation were myopic ( $\beta = 0.08$ ). It would then seem useful to study the money demand dynamics in the context of a more general model.

One may ask why we have not carried out unit root tests as part of this study. Unit root tests are important in examining the stationarity of a time series. Stationarity is a matter of concern in three important areas. First, a crucial question in the ARIMA modelling of a single time series is the number of times the series needs to be first-differenced before an ARIMA model is fitted, as each unit root requires a first-differencing operation. Second, stationarity of regressors is assumed in the derivation of standard inference procedures for regression models. Nonstationary regressors invalidate many standard results and require special treatment. Third, in cointegration analysis, an important question is whether the disturbance term in the cointegrating vector has a unit root.

The relationship between unit root tests and error-correction models relies on the following conditions. If all the dependent variables have a unit root and the error term of the long run model is stationary, then we should use the first order differences of the dependent variables and the lagged error term as regressors, since all of them are stationary.

The theoretical background of the unit root tests is as follows. A series is defined as weakly stationary if it has a finite mean, finite variance and finite covariance, all of which are independent of time. Consider an (AR(1) process

$$Y = m + \rho Y_{-1} + \epsilon$$

where  $m$  and  $\rho$  are parameters and the  $\epsilon$ 's are assumed to be independent and identically distributed with zero and constant variance. The AR(1) process is stationary if  $-1 < \rho < 1$ . If  $\rho=1$  the equation defines a random walk with drift and  $Y$  is then nonstationary. If the process is started at some specific point, the variance of  $Y$  increases steadily with time and the unconditional variance is infinite. If the absolute value of  $\rho$  is greater than one, the series is explosive. Thus the crucial null hypothesis for testing stationarity is that the



absolute value of  $\rho$  should equal one. Since economic series are almost always positively, rather than negatively, correlated the appropriate null hypothesis is

$$H_0: \rho = 1$$

and the test of this hypothesis is a unit root test.

The simplest way to test the null hypothesis is to specify the AR(1) equation as

$$\Delta Y = m + \gamma Y_{-1} + \epsilon$$

where  $\gamma = \rho - 1$ , and  $\Delta Y$  is the first difference of the Y series. Thus the unit root hypothesis is now

$$H_0: \gamma = 0$$

The OLS test produces a t-statistic for  $\hat{\gamma}$ , which may be used to test the significance of  $\gamma$ . The t-statistic, however, cannot be referred to the critical values in the standard 't' table, since under the null hypothesis the Y variable is nonstationary and this table no longer applies. Fuller (1976) derived appropriate limit distributions for the test statistic and Dickey (1976) computed empirical approximations for selected sample sizes. More recently Mackinnon (1990) has implemented a much larger set of replications than those underlying the Dickey-Fuller tables. In addition, he has estimated response surface regressions over these replications. These response surfaces permit the calculation of Dickey-Fuller critical values for any sample size.

The exposition so far has been in terms of the simple AR(1) process. The generalization to higher order processes can be most easily achieved by the use of the lag operator L, where  $LY = Y_{-1}$  etc. The general autoregressive process of order p, AR(p), may be written in the form

$$A(L) Y = m + \epsilon$$

where A(L) is a polynomial of degree p in the lag operator, that is

$$A(L) = 1 - \alpha_1 L - \alpha_2 L^2 - \dots - \alpha_p L^p$$

Replacing L with one gives

$$A(1) = 1 - \alpha_1 - \alpha_2 - \dots - \alpha_p$$

so that A(1) indicates the sum of all the coefficients in the autoregressive scheme. To test whether the pth order process has a unit root, rewrite A(L) as

$$A(L) = (1 - \rho L) B(L)$$

where B(L) is a polynomial of degree (p - 1) in the lag operator. Replacing L with one in this expression gives

$$A(1) = (1 - \rho) B(1)$$

Thus if the process has a unit root then A(1) equals zero, that is, the sum of all the autoregressive coefficients will be zero.

Straightforward estimation of the AR(p) process yields estimates of the  $\alpha_i$  coefficients. These do not permit a simple test of the null hypothesis since these coefficients are all functions of the  $\rho$  parameter. However, a simple test can be obtained by rearranging the autoregression to isolate the  $\rho$  parameter. The appropriate rearrangement is

$$\Delta Y = m - A(1) Y_{-1} + \delta_2 \Delta Y_{-1} + \delta_3 \Delta Y_{-2} + \dots + \delta_p \Delta Y_{-p+1}$$

where the  $\delta$  coefficients are functions of the  $\alpha$ 's. Thus under the null hypothesis of a unit root the coefficient of  $Y_{-1}$  will be zero. If  $\rho < 1$  the coefficient of  $Y_{-1}$  will be negative. This regression is referred to as the Augmented Dickey-Fuller (ADF) regression since the right hand side has been augmented by lagged first differences of Y.

An important result obtained by Fuller is that the limit distribution of the t-statistic on the coefficient of  $Y_{-1}$  is independent of the number of lagged first differences included in the ADF regression. Moreover, standard tables can be used (at least asymptotically) to assess the significance of the  $\delta$  coefficients.

The cointegration tests consists of two steps. The first step is to run a regression for a long run model and record the regression residuals. Step two is to run the unit root test on the residuals. It is important to notice that the unit root tests are based on limit distributions. Thus the tables of critical values that determine test results are valid only asymptotically. The tests may not make much sense when the sample sizes are small. In this case, inspecting the plots of the series and its differences may help checking the stationarity.

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