Economics of Transition
Volume 12 (4) 2004, 653–681

Granger causality of the inflation–growth mirror in accession countries*

Max Gillman* and Anton Nakov**
*Central European University, Department of Economics, Nador utca 9, H1051 Budapest, Hungary. Email: gillman@ceu.hu
**Universitat Pompeu Fabra, Department of Economics and Business, Jaume I Building, Ramon Trias Fargas, 25–27, Barcelona, Spain.

Abstract

The paper presents a model in which the exogenous money supply causes changes in the inflation rate and the output growth rate. While inflation and growth rate changes occur simultaneously, the inflation acts as a tax on the return to human capital and in this sense induces the growth rate decrease. Shifts in the model's credit sector productivity cause shifts in the income velocity of money that can break the otherwise stable relationship between money, inflation, and output growth. Applied to two accession countries, Hungary and Poland, a VAR system is estimated for each that incorporates endogenously determined multiple structural breaks. Results indicate Granger causality positively from money to inflation and negatively from inflation to growth for both Hungary and Poland, as suggested by the model, although there is some feedback to money for Poland. Three structural breaks are found for each country that are linked to changes in velocity trends, and to the breaks found in the other country.

JEL classifications: C22, E31, O42.
Keywords: Granger causality, VAR, transition, inflation, growth, velocity, structural breaks.

*The authors are grateful to Szilard Benk, Tony Braun, Marek Dabrowski, Zsolt Darvas, Stanislaw Gomulka, Jacek Rostowski, Janos Vincze and Myles Wallace for comments, and to the Central European University for research grants.

1. Introduction

Research has investigated both the cause of inflation in transition and the effect of inflation on output and its growth. For example Ross (2000) finds evidence of Granger causality from money to inflation in Slovenia; Nikolic (2000) finds a money–price link in Russia; Hernandez-Cata's (1999) regression analysis of 26 CEE and CIS countries finds that while price decontrol has a one-time effect on the price level, monetary expansion has been the fundamental determinant of inflation; and Sahay and Vegh (1995) find that the market economy relation, whereby money is the main factor in inflation, also applies to transition countries.[1] In terms of inflation and growth, in transitional countries inflation has been found to negatively affect output growth for inflation rates above a threshold rate (Christoffersen and Doyle, 2000), to relate negatively with output growth in four Asian transition economies (including China), and to relate negatively to output growth in 26 transition countries (Lougani and Sheets, 1995). Or as Wyplosz (2000) finds, '*inflation has been found to be incompatible with growth . . . and the choice of the exchange rate regime, another of the early controversies, appears as secondary to the adherence of a strict monetary policy*'.

This paper contributes a VAR analysis of money, prices, and output, from which Granger causality is examined from money to inflation and from inflation to output growth in two accession countries, Hungary and Poland. The empirical investigation is based on an analytic model of money, inflation, and growth in which the income velocity of money is endogenously determined by the relative cost of money versus the cost of credit that is produced in a separate 'banking' sector. In the model, money supply increases cause inflation. Inflation lowers the return to human capital and decreases the growth rate. The empirical results find strong evidence of Granger causality from money to inflation and from inflation to growth for Hungary and for Poland as is suggested by the equilibrium balanced-growth path of the model. Polish results, however, additionally indicate some feedback to the money supply. Several structural breaks are found for both countries. These are explained by shifts in the income velocity of money that 'break' the otherwise stable relation between money, inflation and output growth.

Figures 1–4 present data which suggest a close relation between money, inflation and growth in four transition countries. During certain periods, the inflation data almost mirror the output growth data, a phenomenon we call the 'transition mirror'. The data are for the growth rate of the (CPI) price index, of the real GDP or the industrial production index, and of the money supply (all measured in one year percentage changes) for four EU accession countries. These are two 'first wave' countries, Hungary and Poland, and two 'second wave' countries, Romania

[1] For an example within industrial countries, see Crowder (1998) for evidence of Granger causality from money to inflation for US data.

Figure 1. Hungary: Money growth, inflation, output growth (one-year % changes in money, prices and output)

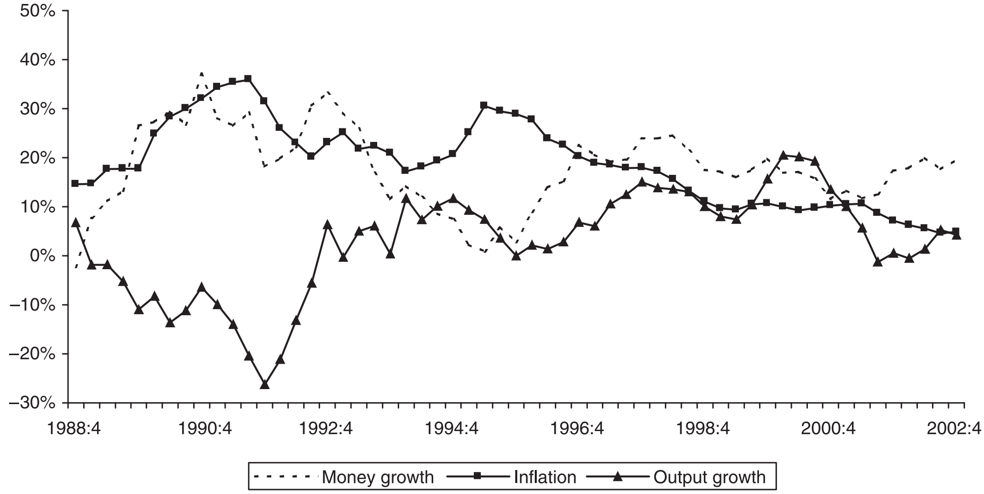
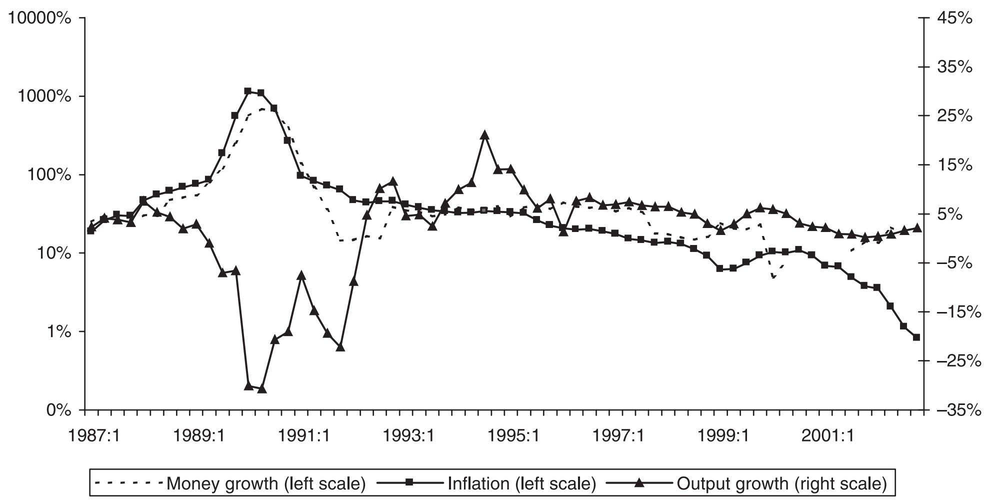


Figure 2. Poland: Money growth, inflation and output growth



and Bulgaria. These countries are chosen on the basis of having the longest IMF compiled data series within the *International Financial Statistics* (2002) on-line database; quarterly data starting in 1987 are provided for Hungary and Poland, and annual data starting in 1986 for Romania and Bulgaria. In all four countries there

Figure 3. Romania: Money growth, inflation, output growth

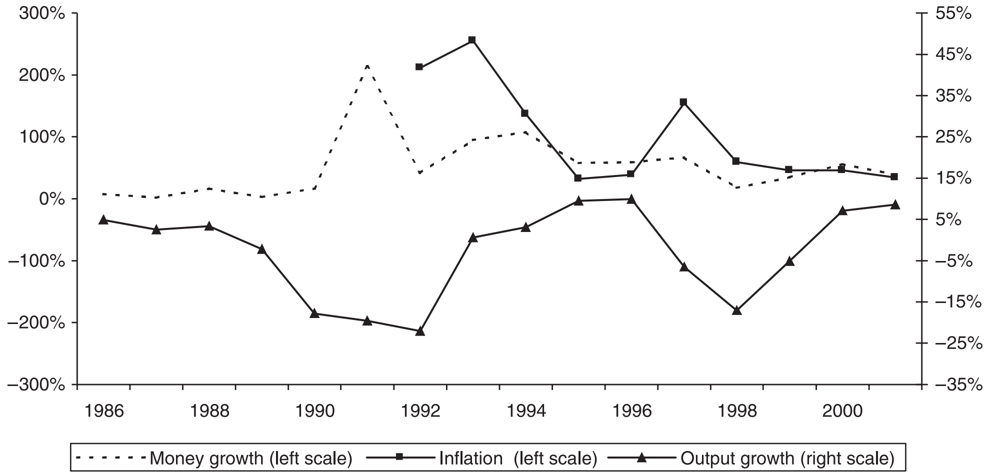
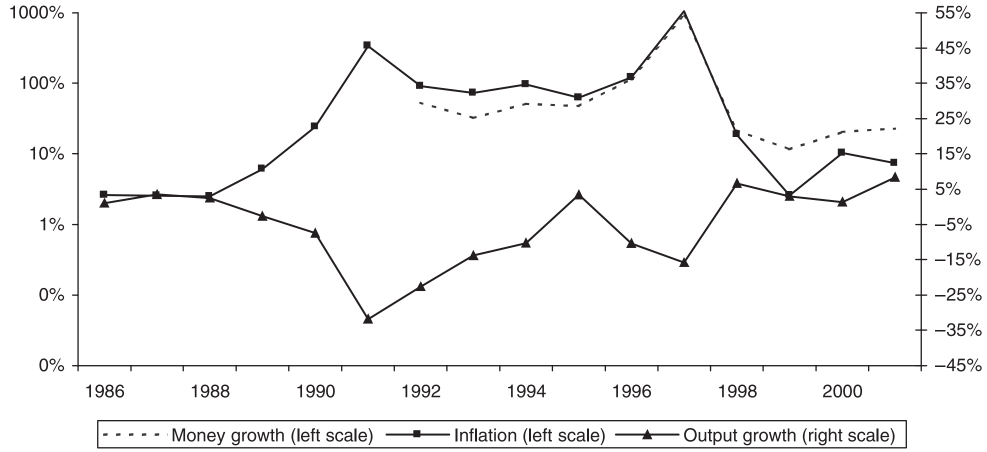


Figure 4. Bulgaria: Money growth, inflation, output growth



is an association of high money growth with inflation. And there is a strong negative correlation between the inflation rate and the growth rate; for example, in Figure 1, Hungary shows this strikingly from 1988 to 1998.

Romania and Bulgaria cannot be tested econometrically because of the paucity of quarterly data. However, some fifteen years of quarterly data exist for Hungary and Poland. Regression analysis, including Granger causality testing, for transition economies requires an important allowance for 'structural breaks' that might affect the relationships among variables; otherwise the regression results can be misleading when assuming parameter constancy over time. In particular, ignoring significant breaks affects both the coefficient estimates and the estimated standard errors, which leads to invalid inference due to model misspecification. To establish the nature of Granger causality in the systems for Hungary and Poland, as based on the monetary endogenous growth economy, we conduct a VAR analysis of money growth, inflation, and output growth. We incorporate endogenously determined multiple structural breaks, estimate the coefficients of the money growth–inflation and inflation–output growth links, and test these relationships for Granger causality. The results offer support for the model and suggest that monetary factors have influenced the course of inflation and growth during the transition stabilization period.

2. The model

The model is an endogenous growth monetary model with growth driven by Lucas type human capital accumulation (Lucas, 1988), and money employed through a modified cash-in-advance (Lucas, 1980) transactions technology that includes credit as an exchange alternative. It is an extension of the Gillman and Kejak (2005) economy, with two main differences. First physical capital is used in goods production, but not in human capital production as in Gillman and Kejak (2005), to make human capital production of a simpler linear form as in Lucas (1988). Second, credit is used not only for consumption, as in Gillman and Kejak (2005), but also for investment. Here it is assumed that the same fraction of both consumption goods and investment goods are bought with credit, where the fraction is determined endogenously within the model in a way similar to Gillman and Kejak (2005). This allows the income velocity of money to be expressed with a closed-form solution, while only the consumption velocity is solved as a closed-form solution in Gillman and Kejak (2005). This extension is important because the income velocity of money plays an important role in explaining structural breaks found in the empirical evidence. Having a closed-form solution of income velocity allows us to explain the breaks with shifts in the parameters affecting velocity.

2.1 Consumer problem

Let the representative consumer's current period utility function be given by the log form:

$$u_t = \ln c_t + \alpha \ln x_t. \tag{1}$$

The consumer allocates time fractionally between working in the goods production sector, l_t, working to produce human capital, l_{ht}, working to produce credit, l_{dt} (we will call this 'banking time') and spending time in leisure, x_t. The allocation of time constraint is:

$$1 = l_t + l_{ht} + l_{dt} + x_t. \tag{2}$$

The consumer accumulates physical capital k_t and rents it to the goods producer, earning a real rental income of $r_t k_t$. Along with the real wage income from effective labour of $w_t l_t h_t$, where h_t is the human capital stock, the consumer spends the income on consumption of goods c_t, on physical capital investment $\dot{k}_t + \delta_k k_t$, and on money stock investment, denoted in nominal terms by $\dot{M}_t$, or in real terms by $\dot{M}_t/P_t$ with P_t denoting the price of the consumption good. A real lump sum transfer from the government adds to the consumer's income, this being the inflation tax proceeds as denoted by V_t/P_t. In real terms we can define $m_t \equiv M_t/P_t$ and $v_t \equiv V_t/P_t$, and can define the inflation rate as $\pi_t = \dot{P}_t/P_t$, and then write $\dot{M}_t/P_t = \dot{m}_t + \pi_t m_t$. This makes the income constraint equal to:

$$r_t k_t + w_t l_t h_t + v_t - c_t - \dot{k}_t - \delta_k k_t - \dot{m}_t - \pi_t m_t = 0. \tag{3}$$

The consumer accumulates human capital, net of the depreciation $\delta_h h_t$, with a function that is linear in the effective time spent in human capital investment, $l_{ht} h_t$. Given the shift parameter $A_h > 0$, and $\delta_h \in [0, 1]$,

$$\dot{h}_t = (1 - \delta_h) h_t + A_h l_{ht} h_t. \tag{4}$$

And the consumer buys the total output, $y_t = c_t + \dot{k}_t + \delta_k k_t$, using either money or credit. The real credit purchases, denoted by d_t, plus the real money purchases m_t, sum to give the total output:

$$m_t + d_t = y_t. \tag{5}$$

The credit is produced by the consumer using the following technology

$$d_t = A_{dt} (l_{dt} h_t)^{\gamma} y_t^{1-\gamma}. \tag{6}$$

This means for example that as the economy progresses along the balanced growth path, with human capital and output growing at the same rate, an increase in the share of goods bought with credit, that is of d_t/y_t, requires an increase in the labour time l_{dt} allocated to credit production. This increase in the share occurs with diminishing returns to labour time, which implies an upward sloping marginal cost curve in producing the share d_t/y_t. The diminishing returns parameter γ determines the convexity of the marginal cost curve, with $\gamma \in (0, 0.5)$ implying a conventional convex marginal cost that rises as the output share rises.

Defining the share of purchases made with money as $a_t \equiv m_t / y_t$, the 'Clower constraint' can be written as

$$m_t = a_t y_t. \tag{7}$$

Substituting Equations (6) and (7) into (5), the credit share is

$$(1 - a_t) = A_{dt}(l_{dt} h_t / y_t)^{\gamma}. \tag{8}$$

Solving for a_t from Equation (8) and substituting this into Equation (7) gives the revised Clower constraint of

$$m_t = [1 - A_{dt}(l_{dt} h_t / y_t)^{\gamma}] y_t. \tag{9}$$

This contains the credit technology. With this as an additional constraint on the optimization problem, like the Clower constraint in monetary economies, the consumer's choice of its banking time will yield a Baumol-type equalization of the marginal cost of money and of credit in equilibrium that in turn determines money demand and velocity (Baumol, 1952). Altogether the consumer's utility maximization is subject to the income (3), human capital (4), and exchange (9) constraints with respect to goods, leisure, goods labour time, human capital time, banking time, and money, human, and physical capital stock levels.[2]

2.2 Goods producer problem

The goods production technology is assumed to be constant returns to scale in effective labour and physical capital. With A_g a shift parameter and $\beta \in (0, 1)$,

$$y_t = A_g (l_t h_t)^{\beta} (k_t)^{1-\beta}. \tag{10}$$

The first-order conditions of the standard profit maximization problem give that

$$w_t = \beta A_g (l_t h_t)^{\beta - 1} (k_t)^{1-\beta}, \tag{11}$$

$$r_t = (1 - \beta) A_g (l_t h_t)^{\beta} (k_t)^{-\beta}. \tag{12}$$

[2] The credit technology is very similar to that in Li (2000), except that Li includes a mechanism designed to induce a liquidity effect. Note that while Li specifies that both labour and capital enter the production of the credit, capital is assumed to be fixed; this is analogous to the assumptions made here, with the fixed capital set equal to one.

2.3 Government money supply

The government supplies new money through lump sum transfers V_t to the consumer so that the money supply evolves as

$$\dot{M}_t = V_t. \tag{13}$$

This occurs at an assumed constant rate σ, where $\dot{M}_t/M_t = V_t/P_t = \sigma$.

2.4 Equilibrium

The consumer's problem can be expressed as a current period Hamiltonian, with maximization with respect to c_t, x_t, l_t, l_{dt}, m_t, k_t, h_t:

$$\begin{aligned} H = {} & e^{-\rho t}(\ln c_t + \alpha \ln x_t) \\ & + \lambda_t(r_t k_t + w_t l_t h_t + v_t - c_t - \dot{k}_t - \delta_k k_t - \dot{m}_t - \pi_t m_t) \\ & + \eta_t[A_h(1 - l_t - l_{dt} - x_t)h_t - \delta_h h_t - \dot{h}_t] \\ & + \mu_t\{m_t - [1 - A_{dt}(l_{dt}h_t/[A_g(l_t h_t)^{\beta}(k_t)^{1-\beta}])^{\gamma}]A_g(l_t h_t)^{\beta}(k_t)^{1-\beta}\} \end{aligned} \tag{14}$$

The equilibrium conditions can be expressed as a reduced set of equations along the balanced-growth path that describe a certain marginal rate of substitution between goods and leisure, an equalization of the return on human capital to the return on physical capital, a balanced-path growth rate denoted by g, an implicit Fisherian equation of the nominal interest rate, denoted by R_t, a closed-form solution for a_t, and the demand for money. These conditions respectively are:

$$\frac{x_t}{\alpha c_t} = \frac{1 + R_t[\gamma + a_t(1-\gamma)]}{w_t h_t} \tag{15}$$

$$\begin{aligned} -\frac{\dot{\mu}_t}{\mu_t} &= A_h(1 - x_t) - \delta_h \\ &= r_t\left\{1 - \frac{a_t R_t}{1 + R_t[\gamma + a_t(1-\gamma)]}\right\} - \delta_k = -\frac{\dot{\lambda}_t}{\lambda_t} \end{aligned} \tag{16}$$

$$\begin{aligned} g &= A_h(1 - x_t) - \delta_h - \rho \\ &= r_t\left\{1 - \frac{a_t R_t}{1 + R_t[\gamma + a_t(1-\gamma)]}\right\} - \delta_k - \rho \end{aligned} \tag{17}$$

$$R_t = r_t - \frac{a_t R_t r_t}{1 + R_t[\gamma + a_t(1-\gamma)]} - \delta_k + \pi_t \tag{18}$$

$$m_t = [1 - A_d^{1/(1-\gamma)}(\gamma R_t/w_t)^{\gamma/(1-\gamma)}]y_t \tag{19}$$

$$a_t = 1 - A_d^{1/(1-\gamma)}(\gamma R_t/w_t)^{\gamma/(1-\gamma)} \tag{20}$$

From these conditions we can fully describe the economics of the model. First note that in the marginal rate of substitution Equation (15), if $a_t = 1$ so that it is a money-only economy, this rate is similar to a Stockman (1981) model extended with human capital. Then the shadow price of goods is 1 plus the nominal interest rate R_t for all purchases. With credit, the exchange cost is less than R_t in general, equal instead to a weighted average of money and credit exchange costs, or $R_t[\gamma + a_t(1-\gamma)]$, which is also equivalent to $a_tR_t + (1-a_t)\gamma R_t$. The average cost of money is R_t and that of credit is γR_t, and the weights are a_t and $(1-a_t)$. When inflation goes up, the nominal interest rate rises, and while a_t falls (see Equation (20)) so that less money and more credit is used, the cost of goods relative to leisure still rises; the agent then substitutes from goods to leisure. This substitution towards leisure causes the return on human capital (Equation (16)) to fall and the growth rate to fall (Equation (17)). There is a subsidiary effect of an increased capital to effective labour ratio in both goods and human capital sectors, a Tobin (1965) effect (see Gillman and Nakov, 2003), until the real return on physical capital falls sufficiently to reestablish equilibrium with the return to human capital.[3] This reallocation of inputs mitigates the fall in the growth rate because of inflation, but is a second-order effect that leaves the growth rate still lower as a result of inflation.

Note that the nominal interest rate R_t is affected by the Stockman (1981) result whereby investment as well as consumption is purchased by money, as shown in the cash in advance constraint. The difference here from Stockman is that only the endogenous fraction a_t of investment is subject to this. This makes the real return to capital contain the inflation tax, and with $a_t = 1$ the exact Stockman result ensues, that $R_t = \frac{r_t}{1+R_t} - \delta_k + \pi_t$. Therefore the equilibrium more generally with $a_t \le 1$ represents an extension of Stockman.

Inflation lowers the growth rate by inducing a lower rate of return to capital.[4] This creates the main link between inflation and growth. The inflation also increases the income velocity of money, y_t/m_t, as seen in Equation (19). Other exogenous factors can also cause a shift in velocity. This is the focus of Gillman and Kejak (2003) who explain changes in the trends in velocity of various US monetary aggregates on the basis of changes in inflation, and in particular changes in the productivity of banking as a result of deregulation. Here bank deregulation is captured by an increase in the productivity parameter A_{dt}. The steady relation

[3] Rapach (2003) finds evidence of a long-run reduction in the real interest rate as caused by inflation in each of 14 industrial countries.

[4] This type of model and its negative effect of inflation on growth is supported empirically by Gillman, Harris and Matyas (2004).

Table 1. Data series

Variable	Notation	Definition	IFS Series Code
Money	m	National Currency	9.434 . . . ZF . . .
Prices	p	Consumer Price Index (1995 = 100)	9.464 . . . ZF . . .
Output	y	Industrial Production or GDP Volume (both 1995 = 100)	9.466 . . . ZF . . .

Notes: The output series for Poland is obtained by splicing Industrial Production, which is available only through 1995, with GDP Volume (IFS code 96499B . . . ZF . . .). We apply 'Census X12' additive seasonal adjustment to the level series.

between money, inflation, and output found in the money demand Equation (19), which might be estimated in a VAR, can be broken with a sudden shift in A_{dt}.

3. Data and empirical methodology

We use quarterly data for Poland from 1986:1 to 2002:4 (68 obs.) and Hungary from 1987:4 to 2002:4 (61 obs.) from the *International Financial Statistics* (2002), IMF. Table 1 describes the data.

Formal testing of the relationships among the variables described in the introduction and backed by our theoretical model takes the following steps. First, we check the order of integration of the series to determine which of them may enter into stable relationships. In these tests we allow for the possibility of structural breaks as opposed to a stochastic trend in the series. Next, we test for cointegration among the I(1) variables using Johansen's maximum likelihood procedure (Johansen and Juselius, 1990). In the absence of cointegration among the I(1) variables, we estimate stationary VAR models with the log-differenced series, allowing for multiple structural breaks in the relationships. We test for Granger-causality (Granger, 1969), show impulse-responses and comment on the variance decompositions.

4. Empirical results

We begin the analysis by examining the univariate statistical properties of the series. We start by applying two standard unit root tests: the Augmented Dickey and Fuller (1979), and the Phillips and Perron (1988), also known as the ADF t and Phillips Z_t tests, respectively. These tests have as null hypothesis that of non-stationarity and the critical values are provided by MacKinnon (1991). As the column labeled 'Standard ADF' of Table 2 shows, on the basis of the conventional ADF test, it appears that the levels of prices and output in Hungary and of output in Poland are I(2) because their first differences appear to have unit roots. Likewise, judging

Table 2. Unit root tests

Variable	ADF t		Phillips Z_t	Order of integration
	Standard	With break		
Hungary				
Money	1.60	−3.34	2.53	I(1)
Money growth	−7.02**	−8.27**	−7.07**	I(0)
Prices	−2.23	−2.87	−3.36	I(1)
Inflation	−2.87	−5.51*	−4.50**	I(0)
Output	−2.12	−4.60	−1.14	I(1)
Output growth	−3.45	−8.55**	−5.92**	I(0)
Poland				
Money	−1.44	−5.50*	−0.85	I(1)
Money growth	−3.93*	−6.12**	−3.41	I(0)
Prices	−1.46	−9.25**	−0.68	I(1)
Inflation	−4.72**	−8.32**	−3.22	I(0)
Output	−2.25	−7.99**	−1.31	I(1)
Output growth	−2.51	−8.63**	−6.41**	I(0)

Notes: *(**) denotes significance at 5%(1%). The 5%(1%) MacKinnon (1991) critical values for the standard ADF t and Phillips Z_t tests including constant and trend are −3.49(−4.12). The 5%(1%) Zivot and Andrews (1992) critical values for the ADF test with break in the level and the trend are –5.08(–5.57). The reported ADF t statistics are for downward-t-chosen autoregressive lag length, while the Phillips Z_t tests use Bartlett kernel with Newey–West lag truncation.

by the standard Phillips Z_t test, money and prices in Poland seem to display I(2) behaviour.

Recent literature has argued that economic time series are unlikely to have such highly non-stationary behaviour. Since Perron (1989), a number of studies have emphasized that rather than possessing a unit root many economic time series may be 'broken-trend stationary'. Perron (1989) showed through a Monte-Carlo experiment that if the magnitude of a discreet shift in the series is significant, standard unit root tests such as ADF t fail to reject the null of non-stationarity even if the series are stationary with a broken trend and *iid* disturbances.

Indeed, pre-testing our series with three standard tests for structural break – the Quandt–Andrews *SupF* test (Quandt, 1960; Andrews, 1993), and the *ExpF* and *AveF* tests of Andrews (1993) and Andrews and Ploberger (1994) using the p-values of Hansen (1997), we find strong evidence of discrete shifts in each of the series for both countries. In light of this finding, we repeat the ADF unit root test, this time allowing for a single structural break in the level and trend of each series. We follow the procedure of Zivot and Andrews (1992), estimating the breakpoint from

Table 3. Cointegration ranks of the systems in levels

Data trend:	None	None	Linear	Linear	Quadratic
Coint. vector:	no c, no trend	c, no trend	c, no trend	c, trend	c, trend
	Number of cointegration relationships chosen by BIC				
Hungary					
3 lags	0	0	0	0	0*
4 lags	0	0	0	0	0
5 lags	1	1	1	0	0
Poland					
3 lags	0	0	0	0	0
4 lags	1	1	0	0	0*
5 lags	1	1	1	1	1

Note: (*) denotes the model which minimizes BIC among the listed models for each country.

the data, searching for the minimum ADF t statistic over all possible break dates. In the column labeled 'ADF with break', Table 2 juxtaposes the results of these tests to the conventional ones.

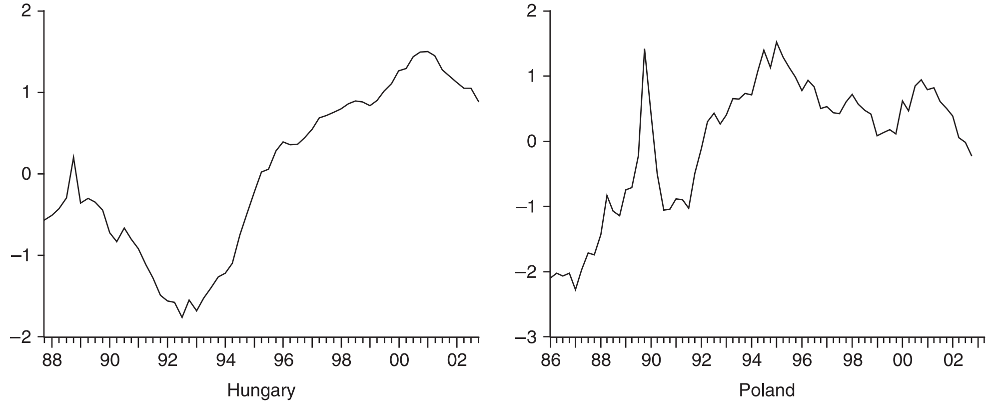
Notice that money growth, inflation and output growth in Hungary all test stationary when applying the ADF test with break, while the levels of money, prices and output for this country all test I(1).[5] These results for Hungary are confirmed by the Phillips Z_t test, which uses a non-parametric approach to controlling for serial correlation.

Interestingly, for Poland the ADF t test with break indicates that all series – levels and growth rates – are broken-trend stationary. However, this (which may be a result of the near-hyperinflation in Poland) is not confirmed by the Phillips Z_t test, which, like the standard ADF test, suggests that the levels are I(1), while the growth rates are stationary.

On the whole we conclude in the last column of Table 2 that while the level series are more likely to be non-stationary, the rate-of-change series are better described as containing discrete shifts rather than stochastic trends. In any case, we are interested in estimating the growth effects of inflation, and establishing stationarity of the growth rates was necessary for correct inference based on a VAR in first differences. Before we proceed with the estimation, following the standard methodology, we test for cointegration among the I(1) levels to see if we should include error-correction terms in the VAR system. Table 3 shows the results from these cointegration tests.

[5] Even though the ADF test statistics 'with break' are uniformly smaller (more negative) than the standard ones by construction, the critical values for the ADF test with break are substantially smaller than those for the standard ADF test.

Figure 5. Non-stationary money velocity in Hungary and Poland (normalized data)



Using the Bayesian information criterion (BIC) as a model selection tool, we fail to find compelling evidence of cointegration among the levels of the three variables. Indeed, while for some of the richer lag specifications BIC indicates the presence of one cointegrating vector, in general the criterion is minimized with fewer lags and under the assumption of no cointegration. For example, among the models for Hungary in Table 3, the model with three lags, a quadratic trend in the data (notice the U-shape of output in Hungary) and no cointegration yields the minimum Bayesian information criterion.

Absence of cointegration is consistent with the notion of an unstable real-money/real income relationship during the transition. In fact, forcing cointegration among the levels of output, prices and money implies imposing a stationary velocity of money. The following Figure 5, which depicts the velocity series for Hungary and Poland (defined as output over real money, normalized into standard deviations from the mean), suggests that the latter is unlikely during the period of transition. Formal unit root tests confirm this conjecture: the ADF t and Phillips Z_t statistics for money velocity in Hungary are -0.92 and -0.67 respectively, while for Poland they are -2.37 and -2.25, pointing to non-stationarity in both cases.

We therefore proceed under the more realistic assumption that the levels of money, prices and output are not cointegrated. This means that we can estimate VAR systems in the stationary growth rates of the three variables without including any error correction terms, a task to which we turn next.

Since the pioneering work of Chow (1960) and Quandt (1960), a number of economists have emphasized the possibility that structural changes may affect the relationships among key economic variables. Such structural breaks are very likely to occur during the transition from a centrally-planned to a market-oriented economy, and may reflect major changes in regulation, the break-up of the CMEA trading system, exchange rate regime shifts, or even changes in the methodology for compiling statistical data, to name a few. In order to account for this possibility, we allow for the existence of multiple structural breaks in the stationary VAR systems. At this point, we consider only partial breaks in the intercepts because allowing for breaks in the slope coefficients too would result in a substantial loss of degrees of freedom given the relatively small sample sizes and the fact that we want to allow many breaks. Nevertheless, our parsimonious approach turns out to provide significant gains in the descriptive power of the models and results in specifications which pass a large number of diagnostic tests.

In general, breaks in the VAR structure need not coincide in time with breaks in the individual series found at the stage of univariate unit root testing. To detect the break dates in the model's relationships, rather than specify them using *a priori* information, we first test each of the VAR equations, applying the full battery of tests developed in Bai and Perron (1998), with issues related to their practical application covered by Bai and Perron (2003). These tests include a *SupF*-type test against a fixed number of breaks, the so-called *double maximum tests*, *UDmax* and *WDmax*, against an unknown number of breaks, a procedure of global minimization of the sum of squared residuals and a sequential procedure using the $SupF(l + 1 \mid l)$ test, as well as the *repartition method* of Bai (1997).[6] To resolve potential discrepancies among the different procedures, we use the Bayesian information criterion for selecting the best model among the models with different numbers of breaks.

While the breaks found in one equation need not coincide with breaks in the other two, adding a potentially insignificant break in a VAR equation is safer than omitting a significant one. Therefore, in the next stage, we augment each equation of the VAR system by the break dummies found in all three equations. This preserves the symmetry of the system and the equivalence between efficient maximum-likelihood and least-squares estimation of an unrestricted VAR.

In this way, we find three structural breaks for Hungary: H-1993:2, H-1996:2 and H-2001:1, and three breaks for Poland: P-1989:3, P-1992:3 and P-1998:3. These are discussed at length and relative to the Section 2 model in the subsequent Section 5.

An alternative way to establish the breakpoints is to estimate them directly from the velocity series, which summarizes the *contemporaneous* relationship among the *levels* of money, prices and output. From this perspective, breaks in the *trend* of velocity correspond to breaks in the *contemporaneous* relationship among the *growth rates* of money, prices and output. Estimating the breaks in the trend of velocity and comparing them with the breaks estimated from the VARs we find surprisingly, that the breakpoints estimated from the velocity series and from the VAR coincide exactly for Hungary but differ for Poland. This is demonstrated in

[6] The procedures, limiting distributions of the estimators and test statistics for all these tests are described in detail in Bai and Perron (1998).

Figure 6. Breaks in the VAR vs Breaks in velocity

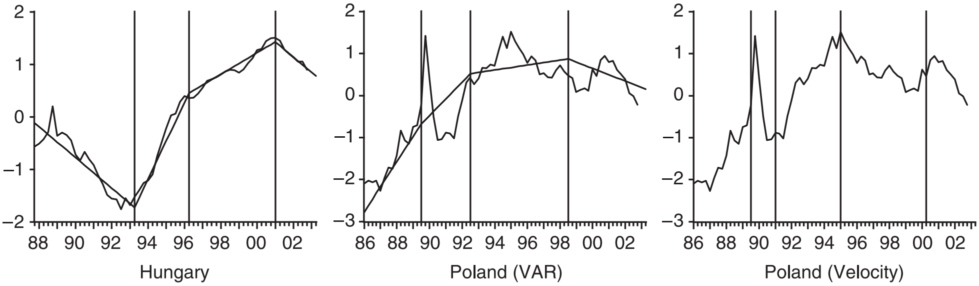


Figure 6, which plots the velocity series together with the break dates shown as vertical lines. In the case of Hungary, estimating the breaks from the VAR yields the same breakpoints as estimating them from a regression of the velocity series on a broken trend, which can be verified visually by the excellent fit of the broken-trend line with the actual velocity series. In the case of Poland, however, we find three breaks in the VAR but four in the velocity series. We attribute this difference to the near-hyperinflation experience in Poland, which makes the contemporaneous relationship (velocity) less stable than the dynamic relationship (captured by the VAR) in which lagged effects play an important role. Since our primary interest is to estimate the dynamic effects of inflation on growth, we choose to work with the VAR-established break dates for Poland.

4.1 Granger causality results

Having determined the break dates T^k, where T^k is the time of the k^{th} structural break, we turn to the estimation of the two break-augmented VAR systems. In particular, we are interested in the cumulative effects and Granger causality of inflation in the output growth equation and of money growth in the inflation equation for each country. Formally, the two VARs that we estimate are:

$$\begin{bmatrix} \Delta m_t \\ \Delta p_t \\ \Delta y_t \end{bmatrix} = \sum_{k=0}^{3} \begin{bmatrix} \alpha_m^k \\ \alpha_p^k \\ \alpha_y^k \end{bmatrix} D_t^k + \sum_{q=1}^{4} \begin{bmatrix} \beta_{mm}^q & \beta_{mp}^q & \beta_{my}^q \\ \beta_{pm}^q & \beta_{pp}^q & \beta_{py}^q \\ \beta_{ym}^q & \beta_{yp}^q & \beta_{yy}^q \end{bmatrix} \begin{bmatrix} \Delta m_{t-q} \\ \Delta p_{t-q} \\ \Delta y_{t-q} \end{bmatrix} + \begin{bmatrix} \varepsilon_t^m \\ \varepsilon_t^p \\ \varepsilon_t^y \end{bmatrix}$$

where $D_t^0 = 1$, $\forall t$, gives the constant, and $D_t^k = \begin{cases} 1 & \text{if } t > T^k \\ 0 & \text{otherwise.} \end{cases}$

Except for the break dummies D_t^k, the above VAR model is standard in the empirical monetary economics literature, as described in Chapter 1 of Walsh (2003). Our tests for Granger causality are in the spirit of Sims (1972). However, because of the issue of non-stationarity, rather than estimating the VAR in levels

Table 4. VAR(4) estimates and Granger causality, Hungary

Equation	Growth (Δy_t) $i \equiv y$	Inflation (Δp_t) $i \equiv p$	Money (Δm_t) $i \equiv m$
$\sum_{q=1}^{4}(\beta_{iy}^{q})$-growth	−0.725**	0.062	−0.462
p-value of F-stat	(0.002)	(0.904)	(0.102)
$\sum_{q=1}^{4}(\beta_{ip}^{q})$-inflation	−1.371**	0.659**	−0.144
p-value of F-stat	(0.000)	(0.000)	(0.795)
$\sum_{q=1}^{4}(\beta_{im}^{q})$-money	0.104	0.139*	−0.402
p-value of F-stat	(0.470)	(0.028)	(0.071)
α_i^0	0.022	0.013	0.079**
p-value of t-stat	(0.223)	(0.147)	(0.000)
α_i^1 (1993:2)	0.077**	0.002	−0.038*
p-value of t-stat	(0.000)	(0.781)	(0.018)
α_i^2 (1996:2)	−0.009	−0.014*	0.038**
p-value of t-stat	(0.419)	(0.012)	(0.002)
α_i^3 (2001:1)	−0.053**	−0.003	−0.010
p-value of t-stat	(0.000)	(0.667)	(0.442)
R^2	0.736	0.799	0.502

Note: *(**) denotes significance at 5%(1%), respectively.

like Sims (1972), we estimate it in the stationary growth rates like Eichenbaum and Singleton (1986). Our analysis is related also to Stock and Watson (1989) and Sims, Stock and Watson (1990) in that we study extensively the stochastic properties of the series to ensure that standard distribution theory can be used to interpret the Granger causality tests.

At the same time, by including multiple structural breaks in our VAR system, we relax the extreme assumption of full parameter constancy over time, made implicitly by other studies of the money-growth link in transition economies (Ross, 2000). In this, our approach is similar to Estrella and Fuhrer (2003) who apply Bai's (1997) test for multiple breaks to a single-equation policy reaction function model, and to Vilasuso (2000) who uses the procedure of Bai and Perron (1998) to establish Granger causality from detrended money to output in the US postwar experience.

We next present the results of the VAR estimation, together with Granger causality tests, in Tables 4, 5, and 6, and plot the dynamic impulse-responses in Figures 7 and 8. Tables 4 and 5 summarize the estimation results for the conventional VAR(4) systems for Hungary and Poland in terms of the cumulative coefficients of each endogenous variable in each equation, together with the probability values of the F tests for joint significance of the estimated coefficients on lags 1 to 4.

Table 5. VAR(4) estimates and Granger causality, Poland

Equation	Growth (Δy_t) $i \equiv y$	Inflation (Δp_t) $i \equiv p$	Money (Δm_t) $i \equiv m$
$\sum_{q=1}^{4} (\beta_{iy}^{q})$-growth	−0.886**	2.654**	1.429**
p-value of F-stat	(0.000)	(0.000)	(0.001)
$\sum_{q=1}^{4} (\beta_{ip}^{q})$-inflation	−0.112**	0.091**	0.330**
p-value of F-stat	(0.000)	(0.000)	(0.000)
$\sum_{q=1}^{4} (\beta_{im}^{q})$-money	−0.178	0.460**	0.387*
p-value of F-stat	(0.181)	(0.002)	(0.024)
α_i^0	0.043**	0.044	0.014
p-value of t-stat	(0.000)	(0.064)	(0.498)
α_i^1 (1989:3)	−0.039	0.235**	0.091*
p-value of t-stat	(0.067)	(0.000)	(0.048)
α_i^2 (1992:3)	0.050*	−0.316**	−0.109*
p-value of t-stat	(0.040)	(0.000)	(0.036)
α_i^3 (1998:3)	−0.035**	0.015	0.006
p-value of t-stat	(0.000)	(0.486)	(0.765)
R^2	0.780	0.906	0.888

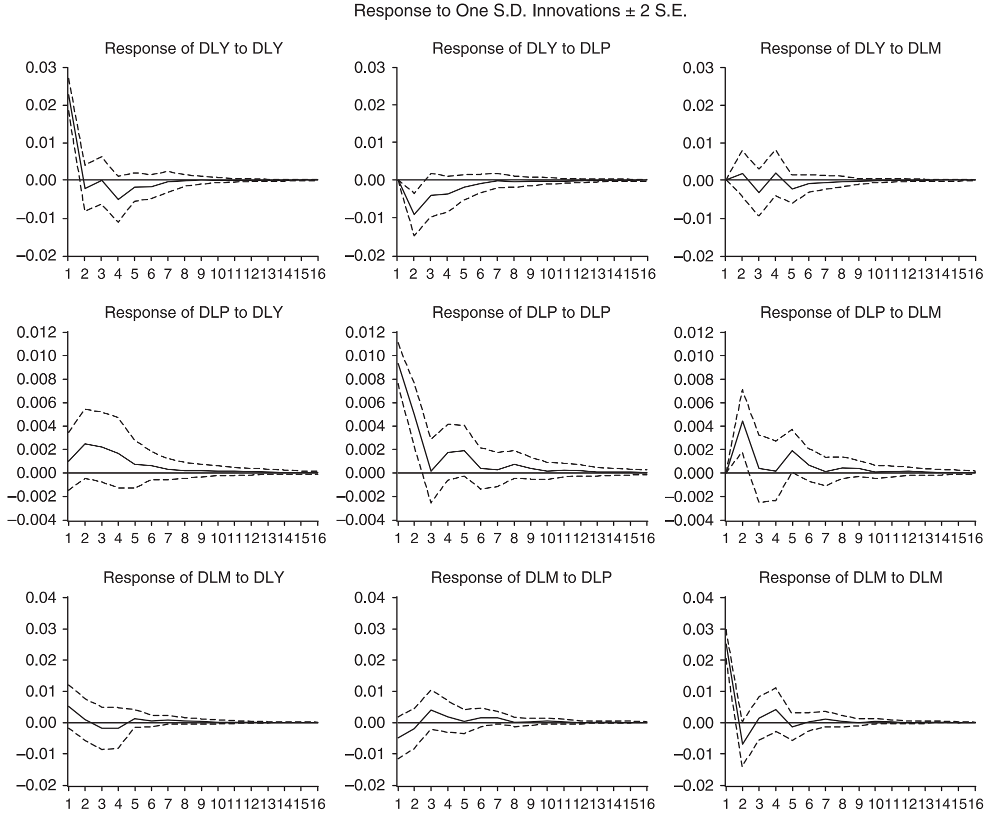
Note: *(**) denotes significance at 5%(1%), respectively.

Table 6. VAR(q) Granger causality tests, q = 3,4,5

Null hypothesis		VAR lag length in quarters		
		[3]	[4]	[5]
Hungary				
Money Growth ↛ Inflation	F-statistic	3.659*	3.044*	1.027
	p-value	0.019	0.028	0.416
Inflation ↛ Growth	F-statistic	5.102**	6.797**	5.711**
	p-value	0.004	0.000	0.000
Growth and Inflation ↛ Money	F-statistic	0.924	1.063	1.273
	p-value	0.487	0.408	0.282
Poland				
Money Growth ↛ Inflation	F-statistic	5.781**	4.908**	2.044
	p-value	0.002	0.002	0.091
Inflation ↛ Growth	F-statistic	11.51**	11.91**	11.64**
	p-value	0.000	0.000	0.000
Growth and Inflation ↛ Money	F-statistic	12.99**	10.61**	8.855**
	p-value	0.000	0.000	0.000

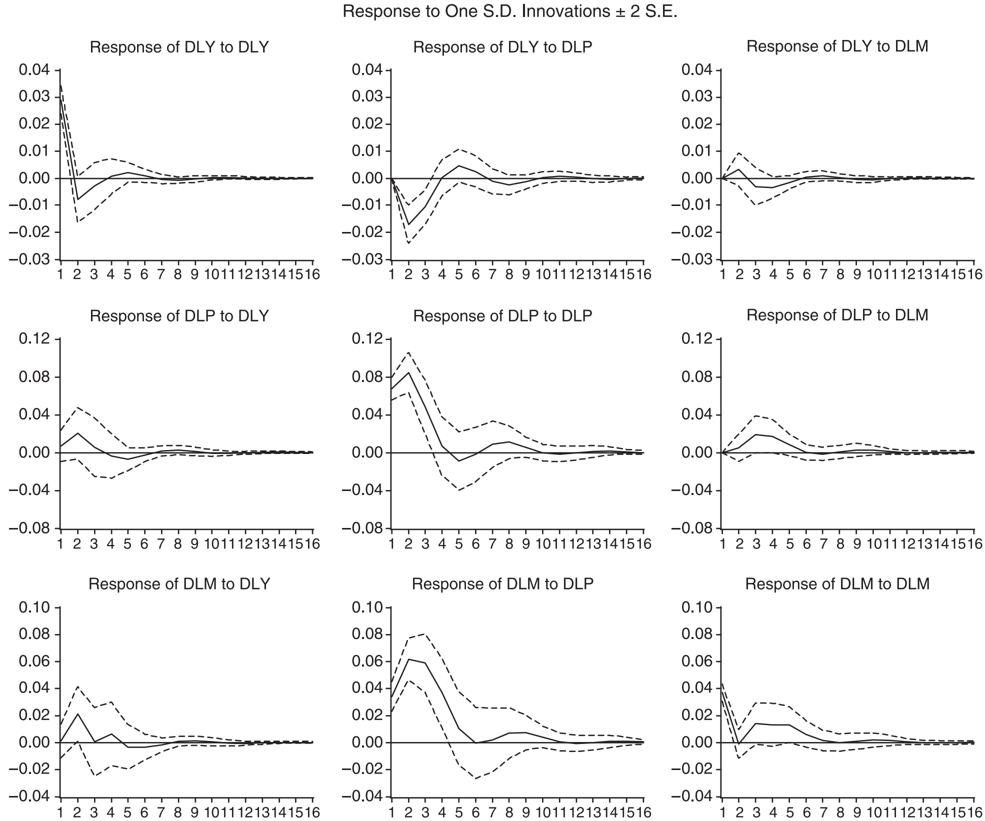
Note: 'No Granger causality' is rejected at 5%(1%) when the p-value is less than 0.05 (0.01), respectively. BIC chooses VAR(3) for Hungary and VAR(4) for Poland.

Figure 7. VAR impulse-responses: Hungary



Thus, for Hungary we find in Table 4 that inflation has a negative cumulative coefficient in the growth equation, estimated at –1.37. Since the *p*-value of the *F* test for joint significance of past inflation in the growth equation is well below 0.05, we discover that there is strong evidence of Granger causality running from inflation to growth. At the same time, money growth has a positive cumulative coefficient in the inflation equation, estimated at 0.14. In addition, the *F* test of joint significance of past money growth on inflation rejects Granger non-causality at the standard 5 percent level. On the other hand, neither past output growth, nor past inflation are significant at 5 percent in explaining money growth, supporting the hypothesis of exogeneity of the growth rate of money supply in the system including the three variables. Furthermore, we find no evidence that output growth Granger-causes inflation, or that money growth directly Granger-causes output growth in the Hungarian transition experience.

Figure 8. VAR impulse-responses: Poland



Turning to Poland, Table 5 shows that in this country, too, money growth Granger-causes inflation, with an estimated cumulative coefficient of 0.46 for the standard VAR(4). As in Hungary, inflation in Poland also affects output growth negatively, with an estimated cumulative coefficient of −0.11. However, in contrast to Hungary, in the case of Poland we reject money growth exogeneity in light of evidence of Granger causality running also from output growth and inflation to money growth, possibly reflecting the reaction of the monetary authority during the period of hyperinflation.

The above results are not very sensitive to changes in the VAR lag specification. For example, the estimated cumulative coefficient of inflation on growth in Hungary lies between –1.13 for VAR(3) and –1.57 for VAR(5) and Table 6 shows that in each case the *F* test finds strong evidence of Granger causality from inflation to growth.

Likewise, the estimated cumulative coefficient of inflation on growth in Poland is in the range between -0.11 and -0.27 and Granger non-causality is rejected strongly with three and five lags as well. Note that our choice of lag length is supported formally by the Bayesian information criterion, which selects three lags in Hungary and four lags in Poland. Moreover, these parsimonious specifications pass a number of diagnostic tests such as (vector) Portmanteau, error autocorrelation, normality and heteroskedasticity tests.

4.2 Impulse responses and variance decomposition

Finally, in Figures 7 and 8 we show the dynamic impulse-responses of the VAR systems for Hungary and Poland, using the orthogonal Cholesky decomposition to identify the shocks. Since we have found that cross-correlation among the VAR residuals is quite small, the resulting impulse-responses are not sensitive to the ordering of the variables in the systems. The responses to a shock in any of the variables die out in less than four years, which is consistent with our conjecture of stationarity of the system. In particular, the negative response of growth to a single one-standard-deviation shock in inflation lasts for about 8–10 quarters in Hungary and Poland. Similarly, the positive response of inflation to a single one-standard-deviation shock in money growth dies out in about two years in both countries. Finally, observe in the third row that money growth essentially does not respond to either inflation or output growth shocks in Hungary but it does respond to inflation shocks in Poland.

In terms of variance decomposition, in Hungary a shock to money growth accounts for about 15 percent of the variation in inflation in four years' time, while an inflation innovation contributes up to 19 percent of the output growth variation over the same period. In Poland, the contribution of a money growth shock to inflation variation in four years' time is about 26 percent, while the share of variance in output growth due to a four-years old inflation innovation is about 33 percent.

To summarize, we find Granger causality with a positive effect running from the money growth rate to inflation, and Granger causality with a negative effect running from inflation to output growth in the transition experience of both Hungary and Poland. We take this econometric evidence as strong support for the existence of a positive link from money growth to inflation and a negative link from inflation to output growth, in line with the theoretical prediction of Section 2. We attribute the feedback from output to money and to inflation in Poland to the near-hyperinflation experience of this country.

5. Structural breaks in Hungary and Poland

The breaks reported in Section 4 are found to correspond to changes in the velocity trends. And surprisingly there are three similar events that happened in each country at different times. This gives a pairwise explanation with three events.

5.1 Hungary

Before detailing our explanation of the breaks consider the contrast between the velocity graphs in Figures 5 and 6 with the inflation graphs in Figures 1 and 2. Consider Hungary first. The inflation rate rose and then fell from 1988 to 1993 roughly similar to the rise and fall in velocity over the same period. In 1993 the inflation rate starting rising again, but only for some six quarters before resuming a permanent trend downwards. In contrast the velocity of money graph shows a rapid climb that continues right up until 2001. This upward trend shows a break to a less steep trend in 1996, but it remains a period of strong velocity increases until 2001. This is completely at odds with the movement of the inflation rate, which by itself would induce a downward trend in velocity. A candidate explanation from the Section 2 model is that the productivity of banking shifted upwards because of deregulation; this would cause velocity to increase. Indeed there were major financial sector deregulations that occurred in 1993 and in 1996.

A major bank refinancing of bad loans started in late 1992 and continued until 2000 with a cost of approximately 13 percent of GDP. The consolidation and restructuring of the bank sector took place in stages. This included ‘cleaning’ the portfolios of the banking sector, where in the second half of 1993 certain large state-owned firms had their bad debts taken off the books of banks in exchange for government bonds. A dramatic drop in non-performing loans as a percentage of total loans took place: from 30 percent in 1993 to 20 percent in 1994, and down to close to 10 percent in 1995, with gradual decreases in all but one year thereafter to below 5 percent by 2000. This ushered in a new era of creating acceptable capital adequacy ratios that enabled banks to move towards the international standards of a competitively functioning bank sector.

Another major event occurred at the end of 1995, the privatization of the bank sector. This began with the selling of six state-owned banks, with a 31 percent market share, to foreign banks. The largest Hungarian bank, the NSB with a 29 percent market share before privatization, was privatized through the stock exchange. State ownership continued to drop until by 1997 it was only 20 percent of the banking sector’s capital. Szapary (2001) details these changes in the Hungarian banking sector.

The two Hungarian episodes of major bank deregulation each represent shifts upward in bank sector productivity, A_{dt} in the theoretical model, that cause upwards shifts in the velocity. Thus we explain the first two Hungarian shifts in H-1993:3 and H-1996:2 in this way.

A different radical change occurred in the Hungarian banking sector with the passage of a new Central Bank Act in July 2001. This was a reform of the Hungarian central bank, the National Bank of Hungary, via a new charter that instituted inflation rate targeting instead of the previous practice of exchange rate targeting. This aimed to reduce the variance of the inflation rate and to lower its level down towards the 0–2 percent international norm among central banks that target the inflation rate. A dramatically lower expected variance in the inflation rate is outside of the theoretical framework of the Section 2 model, which is deterministic. In a stylized way within the model, this lower variance can act as a negative shock to the productivity of the private bank sector in our model, in that banks would no longer play as large a role in allowing agents to avoid fluctuations in the inflation tax. The inflation rate did begin falling in 2001, but it was only a gradual fall, while the velocity abruptly began trending downwards in a way that cannot be explained only by the change in the average inflation rate. A dramatic shift down in the expected variance, acting as a decrease in the bank productivity shift factor, offers an explanation of this velocity shift, and thus we explain the shift H-2001:1 in this way.

5.2 Poland

Velocity versus the inflation trends in Poland were similar to those of Hungary with respect to the times at which the empirically identified structural breaks occurred, although as Figure 6 shows there are some differences related to the hyperinflation. At the end of 1989, the inflation rate peaked and began falling rapidly, and trended downwards mostly from then onwards. This would suggest that velocity would also fall rapidly and then trend downwards as based on an explanation using only the inflation rate. Velocity did initially fall as hyperinflation receded, but it then levelled off and began rising in 1991. As inflation continued steadily downwards in the 1992 to 1994 period, the velocity again acted in the opposite direction as expected from the inflation data alone, with a further steady shift upwards from 1992 to 1995. After that, velocity trended down as did the inflation rate, until the end of 1998. Then Poland experienced an initial increase in the inflation rate for almost two years, before inflation finally steadily moved down towards one percent. Velocity shifted up as did the inflation rate in 1999–2000 and then began a sharp downwards movement.

The divergences of velocity trends from the inflation rate path, near to the break periods of P-1989:3 and P-1992:3, is markedly similar to the experience in Hungary, near to the breaks of H-1993:2 and H-1996:2. The Section 2 model suggests that a candidate explanation for these divergences is shifts in the banking sector productivity parameter. And the following description supports the conclusion that the empirically identified shifts occur largely in line with banking sector deregulations.

On January 1 1989, Poland passed the Banking Act and the National Bank of Poland Act that separated from the central bank nine commercial banks, thereby creating the ‘two-tier’ model of banking. Also legislation was introduced in 1989 that allowed individuals, including foreigners, to form new banks as limited stock companies, with some 70 licences issued from 1989 to 1991. This deregulation continued with privatization of the Export Development Bank in October 1991, and with the nine state-owned commercial banks transformed into limited stock companies. These events effectuated a massive deregulation of banking that started in 1989.

Another banking act was passed in March 1992 that allowed for standard enforcement of capital adequacy and loss provisions. Also a programme with the IMF and World Bank was established for 'twinning' whereby Western banking methods were introduced into the Polish bank sector. In November 1992, the central bank required banks to provision fully against loans, and in March 1993 an Enterprise and Bank Restructuring Program was begun to recapitalize bad loans. This involved a one-time recapitalization of $520 million of the bank sector. Together these regulatory changes resulted in a recapitalization of the bad loans of the banking system, in a fashion similar to what happened in Hungary. Gray and Holle (1996) and Mondschean and Opiela (1997) provide extensive details of these two different types of Polish bank restructurings that began in 1989 and 1992.

August 1997 brought a new central bank independence act, the National Bank of Poland Act, that established inflation rate targeting, or 'price stability', as its main objective. In November 1998 the complimentary Public Finances Act was passed that prohibits funding of the public sector by the central bank. Initially this could be considered as acting as a decrease in the expected variance of the inflation rate that, in the terms of our model, might be described as a shift down in the productivity of banking in avoiding inflation tax. As with Hungary, this is how we explain the break here, but it is less clearly visible for Poland in that the new inflation targeting policy in Poland appears to have been less credible initially since the inflation rate rose at first. This makes it less discernible to what extent velocity may have risen by less and then fallen by more, as a result of the new policy, than could be readily explained by inflation changes alone.

5.3 Pairwise breaks in Hungary and Poland

To summarize:

1. The H-1993:3 and P-1992:3 breaks correspond to a massive refinancing of the bad loans in the state-owned banks. This involved restructuring and consolidation of the banks, and allowed the banks to go forward on a more internationally competitive basis after that point. This acted as a shift up in the productivity of the banking sector that pressured velocity upwards even though inflation rates were increasing.
2. The H-1996:2 and P-1989:3 breaks correspond to major bank privatization laws. These also pressured velocity upwards because of a shift upwards in bank productivity.
3. The H-2001:1 and P-1998:3 breaks correspond to new national bank acts in which inflation targeting was adopted by law. This can be thought of as bringing about a significant change in the expected variance and mean of the inflation rate. Such a reduction in inflation uncertainty can act like a shift down in the banks productivity in producing exchange credit, or other instruments that can be used to avoid the inflation tax, since the value of this avoidance becomes lower as the variance of inflation falls.

6. Discussion

Two points are especially worth discussing further. One issue is whether the model is appropriate for analysing periods of hyperinflation, and a second is whether other factors unrelated to credit sector productivity may be the cause of the structural shifts found in the empirical results. The Cagan (1956) model of money explains hyperinflation as part of a stable money demand function. Since it is not derived from a dynamic general equilibrium model, we cannot really say if it is a long-run or short-run model. But we can see that others have found this model a reasonable description of long-run stable money behaviour. For example Mark and Sul (2002) provide strong evidence of a stable Cagan (1956) money demand for an international panel dataset in which they find a cointegrated money demand function with an income elasticity of 1.08 and a semi-interest elasticity of -0.02.

The model presented here provides a general equilibrium version of a model that is similar to the Cagan (1956) model. In particular, as Gillman and Kejak (2002) show through calibration of a closely related model, as the nominal interest rate rises the magnitude of the interest elasticity rises nearly in proportion to it. The calibrated semi-interest elasticity is nearly constant, depending on the specifics of the calibration. One difference relative to the Cagan (1956) model concerns the paradox, pointed out by Cagan (1956) and Lucas (2000): Cagan (1956) finds a seigniorage-revenue maximizing rate of inflation at $R^* = -1/b$, where b is the estimated constant semi-interest elasticity, while the hyperinflation rates actually observed were clearly above this level.

The paradox is offered a resolution by Marcet and Nicolini (2003). They assume a Cagan-type model of money demand, rationalized by an overlapping generations economy, but suggest a learning process whereby agents can shift their expectations from an adaptive process that is a simple average of past inflation rates to one that more fully understands the onset of a hyperinflation. This 'tracking' model weighs the most recent inflation rates most heavily, with the result that the seigniorage path continues to rise slightly even as the inflation rate rises exponentially. Such a gradually rising seigniorage is also found in Eckstein and Leiderman (1992), in their Sidrauski (1967)-based explanation of Israeli seigniorage.

In the model presented here, the magnitude of the interest elasticity of money starts at zero and rises steadily as the inflation rate rises. But it does not reach one in magnitude, the revenue maximizing point, until very high levels, typically hyperinflation levels depending on the calibration. Thus the model of this paper, like Marcet and Nicolini (2003), does explain a stable money process during hyperinflation, *when the hyperinflation is expected*. And like Eckstein and Leiderman (1992), it is consistent with a seigniorage that approaches a levelling off as the inflation rate rises, even up to hyperinflation rates of inflation. However, it does not explain unexpected surges in inflation.

The paper is potentially able to explain the full Polish experience given that the hyperinflation was expected, and this is possible given the budget deficits being experienced at the time. However the result that Granger causality evidence was also found from output growth and inflation to money for Poland indicates some feedback that may have been a result of the hyperinflation. In particular, if some of the hyperinflation experience were not fully anticipated, possible Phillips curve effects may arise initially that can conceivably lead to such feedback. For example in Poland there may have been an initially delayed shift in the 'tracking' expectations regime that Marcet and Nicolini (2003) describe.

It may also be possible that the breaks in velocity were caused by other factors than shifts in the productivity of the finance sector, A_d. For example, keeping in mind the typical sources of shocks found in the real business cycle literature, the total productivity factor of goods output A_g may have been a source of structural breaks, or even the productivity factor for the human capital production sector, A_h. To consider what effects these may have had consider the equilibrium conditions of the model.

Equation (20) gives the solution for the inverse of the income velocity of real money demand, which is defined by the three variables entering the VAR: the money stock, aggregate price and real output. Should there be a productivity shock through A_g, then inverse velocity is affected through the real wage and real interest rates of Equations (11) and (12). These enter Equation (20) through the ratio of the nominal interest rate to the real wage, R/w. Using the Fisher equation of interest rates, this ratio can be written as $(r + \pi)/w$.[7] A shift in A_d would effectively cancel out for the r/w part of this, leaving it to affect only π/w in the equation; and in this way a positive shock could decrease velocity $(1/a)$. Such an effect is possible but difficult to uncover because of a lack of evidence on total factor productivity in Hungary and Poland.

A shock from the A_h factor for the productivity of human capital investment cannot easily be tracked in Equation (20), as it would enter only indirectly through the capital to effective labour ratios that enter Equations (11) and (12). And such evidence on A_h would seemingly be even more difficult to uncover than for A_g.

Thus while other factors may be behind the VAR structural breaks, corroborating evidence is presented for the shift being from the A_d factor. Further, other studies have found structural breaks that are not inconsistent with this explanation. Using similar Bai and Perron (1998) techniques as in this paper to find structural breaks in inflation series, Benati and Kapetanios (2002) for example find breaks in New Zealand in 1989, in Canada in 1991 and in the UK in 1991 which are interpreted as being due to those countries' adoption of inflation rate targeting. And such inflation rate targeting was described as being related to one of the structural breaks for both Hungary and Poland. Also with the same Bai and Perron (1998) techniques, Vilasuso (2000) examines a money and output VAR for the US from

[7] The Fisher equation can be derived formally within the model by including nominal bonds, but this is suppressed to economize on notation.

1960 to 1997 and finds causality from money to output with two structural breaks, in 1984 and 1991. Benk, Gillman and Kejak (2004) identify business cycle shocks from the credit sector for US data in the 1983–85 and the 1990–92 periods that they associate with changes that followed new banking laws. In particular these were the Garn – St. Germain Act of 1982 that significantly deregulated the banking sector and the Financial Institutions Reform, Recovery and Enforcement Act of 1989 that was designed to clean up the bad loans of the savings and loans industry. A bank deregulation and bad loan clean-up are also associated with the two other structural breaks found for both Hungary and Poland in this paper.

7. Conclusion

The paper presents a dynamic general equilibrium monetary economy with a closed form solution for the income velocity of real money demand. The economy includes the production of credit that enables the consumer to avoid inflation tax. This formulation makes money demand and its velocity depend on structural parameters of credit technology rather than utility parameters as in the Sidrauski (1967) approach or the Lucas and Stokey (1987) approach, or transaction cost parameters in shopping time economies. Unlike these other approaches, here productivity shifts in the production of credit can shift the velocity of money demand. The model also shows how the money supply side of the money market affects the economy through its imposition of inflation tax. This implicit tax reduces the return on human capital and the economy's growth rate. And when there are changes on the money demand side, from changes in productivity in the credit sector, the effect of inflation tax on growth is altered.

Empirical models of the effect of money on inflation and of inflation on growth can as a result be affected by shifts in velocity. This appears to be reflected in the results presented here on structural breaks in the VAR systems. These breaks are explained in terms of shifts in velocity caused by major changes in banking laws.

With the structural breaks, evidence supports Granger causality from money growth to inflation and from inflation to output growth for both Hungary and Poland, leading accession countries. Such evidence provides support for the endogenous growth model in which increases in the money supply growth rate cause the inflation rate to go up, which in turn acts as a tax that causes the output growth rate to fall. For Poland there is also Granger causality of output growth and of inflation on money, which is not explained by the model. A difference between the two countries is that Poland experienced hyperinflation while Hungary did not. Some of the hyperinflation in Poland may have been unanticipated and part of a feedback process between money and output.

The strong results provide support for a monetary-type explanation for part of the transitional recessions experienced in these countries. This may warrant investigating such possibilities in other transition countries, especially as the data become more available; data limitations currently constrain such a broader inquiry. The thesis is meant as an addition to the other hypotheses in the literature that attempt to explain the transitional recessions, as well as indicating the potential importance for developing countries to have low, stationary, inflation rates.

References

Andrews, D. W. K. (1993). 'Tests for parameter instability and structural change with unknown change point', *Econometrica*, 61, pp. 821–56.

Andrews, D. W. K. and Ploberger, W. (1994). 'Optimal tests when a nuisance parameter is present only under the alternative', *Econometrica*, 62, pp. 1383–1414.

Bai, J. (1997). 'Estimating multiple breaks one at a time', *Econometric Theory*, 13(3), pp. 315–52.

Bai, J. and Perron, P. (1998). 'Estimating and testing linear models with multiple structural changes', *Econometrica*, 66, pp. 47–78.

Bai, J. and Perron, P. (2003). 'Computation and analysis of multiple structural change models', *Journal of Applied Econometrics*, 18, pp. 1–22.

Baumol, W. (1952). 'The transactions demand for cash: An inventory – theoretic approach', *Quarterly Journal of Economics*, 66, pp. 545–66.

Benati, L. and Kapetanios, G. (2002). 'Structural breaks in inflation dynamics', *Bank of England manuscript*.

Benk, S., Gillman, M. and Kejak, M. (2004). 'Credit shocks in a monetary business cycle', *working paper 7/2004*, Dept. of Economics, Central European University: Budapest.

Cagan, P. (1956). 'The monetary dynamics of hyperinflation', in Friedman, M. (ed.), *Studies in the Quantity Theory of Money*, The University of Chicago Press, Chicago, pp. 25–120.

Chow, G. (1960). 'Tests of equality between sets of coefficients in two linear regressions', *Econometrica*, 28, pp. 591–605.

Christoffersen, P. and Doyle, P. (2000). 'From inflation to growth: Eight years of transition', *Economics of Transition*, 8(2), pp. 421–51.

Crowder, W. J. (1998). 'The long-run link between money growth and inflation', *Economic Inquiry*, 36(2), pp. 229–43.

Dickey, D. and Fuller, W. (1979). 'Distribution of the estimators for autoregressive time series with a unit root', *Journal of the American Statistical Association*, 74, pp. 427–31.

Eckstein, Z. and Leiderman, L. (1992). 'Seigniorage and the welfare cost of inflation: Evidence from an intertemporal model of money and consumption', *Journal of Monetary Economics*, 29(3), pp. 389–410.

Eichenbaum, M. and Singleton, K. J. (1986). *Do Equilibrium Real Business Cycle Theories Explain Postwar US Business Cycles?, NBER Macroeconomics Annual, 1986*, Cambridge, MA: MIT Press, pp. 91–135.

Estrella, A. and Fuhrer, J. C. (2003). 'Monetary policy shifts and the stability of monetary policy models', *Review of Economics and Statistics*, 85(1), pp. 94–104.

Gillman, M., Harris, M. and Matyas, L. (2004). 'Inflation and growth: Explaining a negative effect', *Empirical Economics*, 29(1), pp. 149–67.

Gillman, M. and Kejak, M. (2002). 'Modeling the effect of inflation: Growth, levels, and Tobin', in Levine, D. (ed.), *Proceedings of the 2002 North American Summer Meetings of the Econometric Society: Money*. http://www.dklevine.com/proceedings/money.htm.

Gillman, M. and Kejak, M. (2003). 'The demand for bank reserves and other monetary aggregates', *Economic Inquiry*, 42(3), pp. 518–33.

Gillman, M. and Kejak, M. (2005). 'Inflation and balanced-path growth with alternative payment mechanisms', *Economic Journal*, forthcoming. See *discussion paper 2004/2*, Hungarian Academy of Sciences, Institute of Economics, Budapest.

Gillman, M. and Nakov, A. (2003). 'A revised Tobin effect from inflation: Relative input price and capital ratio realignments, US and UK, 1959–1999', *Economica*, 70(279), pp. 439–50.

Granger, C. (1969). 'Investigating causal relations by econometric models and cross-spectral methods', *Econometrica*, 37, pp. 424–38.

Gray, C. W. and Holle, A. (1996). 'Bank-led restructuring in Poland', *World Bank Policy Research Working Paper, 1650*, Washington, DC: The World Bank.

Hansen, B. (1997). 'Approximate asymptotic *p* values for structural change tests', *Journal of Business and Economic Statistics*, (1), pp. 60–67.

Hernandez-Cata, E. (1999). 'Price liberalization, money growth, and inflation during the transition to a market economy', *International Monetary Fund Working Paper 99/76*, Washington, DC: IMF.

International Financial Statistics (2002). Database. Washington, DC: International Monetary Fund.

Johansen, S. and Juselius, K. (1990). 'Maximum likelihood estimation and inferences on cointergration-with applications to the demand for money', *Oxford Bulletin of Economics and Statistics*, 52, pp. 169–210.

Li, V. (2000). 'Household credit and the monetary transmission mechanism', *Journal of Money, Credit and Banking*, 32(3), pp. 335–56.

Loungani, P. and Sheets, N. (1995). 'Central bank independence, inflation and growth in transition economies', *International Finance Discussion Papers No. 519*, Washington, DC, by Board of Governors of the Federal Reserve System.

Lucas, Jr., R. E. (1980). 'Equilibrium in a pure currency economy', *Economic Inquiry*, 43, pp. 203–20.

Lucas, Jr., R. E. (1988). 'On the mechanics of economic development', *Journal of Monetary Economics*, 22, pp. 3–42.

Lucas, Jr., R. E. (2000). 'Inflation and welfare', *Econometrica*, 68(2), pp. 247–75.

Lucas, Jr., R. E. and Stokey, N. L. (1987). 'Money and interest in a cash-in-advance economy', *Econometrica*, 55, pp. 491–513.

MacKinnon, J. (1991). *Critical Values for Cointegration Tests*, Oxford: Oxford University Press, chapter 13.

Marcet, A. and Nicolini, J. (2003). 'Recurrent hyperinflation and learning', *American Economic Review*, 93(5), pp. 1476–98.

Mark, N. and Sul, D. (2002). 'Cointegration vector estimation by panel DOLS and long run money demand', *NBER Technical Working Paper 287*, Cambridge, MA: NBER.

Mondschean, T. S. and Opiela, T. P. (1997). 'Banking reform in a transition economy: The case of Poland', *Federal Reserve Bank of Chicago Economic Perspectives*, pp. 16–32.

Nikolic, M. (2000). 'Money growth-inflation relationship in postcommunist Russia', *Journal of Comparative Economics*, 28(1), pp. 108–33.

Perron, P. (1989). 'The great crash, the oil price shock and the unit root hypothesis', *Econometrica*, 57, pp. 1361–401.

Phillips, P. and Perron, P. (1988). 'Testing for a unit root in time series regression', *Biometrika*, 75, pp. 335–46.

Quandt, R. (1960). 'Tests of the hypothesis that a linear regression system obeys two separate regimes', *Journal of the American Statistical Association*, 55, pp. 324–30.
Rapach, D. (2003). 'International evidence on the long-run impact of inflation', *Journal of Money, Credit and Banking*, 35(1), pp. 23–48.
Ross, K. (2000). 'Post-stabilization inflation dynamics in Slovenia', *Applied Economics*, 32(2), pp. 135–49.
Sahay, R. and Vegh, C. A. (1995). 'Inflation and stabilization in transition economies: A comparison with market economies', *International Monetary Fund Working Paper No. 95/8*, Washington, DC: IMF.
Sidrauski, M. (1967). 'Rational choice and patterns of growth in a monetary economy', *American Economic Review*, 57(2), pp. 534–44. Papers and Proceedings of the Seventy-Ninth Annual Meeting of the American Economic Association.
Sims, C. A. (1972). 'Money, income and causality', *American Economic Review*, 62(4), pp. 540–42.
Sims, C., Stock, J. H. and Watson, M. W. (1990). 'Inference in linear time series models with some unit roots', *Econometrica*, 58(1), pp. 113–44.
Stock, J. H. and Watson, M. W. (1989). 'Interpreting the evidence on money-income causality', *Journal of Econometrics*, 40(1), pp. 161–81.
Stockman, A. C. (1981). 'Anticipated inflation and the capital stock in a cash-in-advance economy', *Journal of Monetary Economics*, 8(3), pp. 387–93.
Szapary, G. (2001). 'Banking sector reform in ungary: Lessons learned, current trends and prospects', *National Bank of Hungary Working Paper 2001/5.*
Tobin, J. (1965). 'Money and economic growth', *Econometrica*, 33(4), pp. 671–84.
Vilasuso, J. (2000). 'Trend breaks in money growth and the money-output relation in the U.S.', *Oxford Bulletin of Economices and Statistics*, 62(1), pp. 53–60.
Walsh, C. (2003). *Monetary Theory and Policy*, second edition, Cambridge, MA: MIT Press.
Wyplosz, C. (2000). 'Ten years of transformation: Macroeconomic lessons', *Center for Economic Policy Research Discussion Paper 2254*, London: CEPR.
Zivot, E. and Andrews, D. (1992). 'Further evidence on the great crash, the oil-price shock, and the unit-root hypothesis', *Journal of Business and Economic Statistics*, 10, pp. 251–70.