Long-Run Links Among Money, Prices, and Output: World-Wide Evidence

Helmut Herwartz
(Humboldt-Universität zu Berlin)

Hans-Eggert Reimers
(Hochschule Wismar)

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Deutsche Bundesbank, Wilhelm-Epstein-Strasse 14, 60431 Frankfurt am Main,
P.O.B. 10 06 02, 60006 Frankfurt am Main, Federal Republic of Germany

Telephone (0 69) 95 66-1
Telex within Germany 4 1 227, telex from abroad 4 14 431, fax (0 69) 5 60 10 71
Internet: http://www.bundesbank.de

Please address all orders in writing to: Deutsche Bundesbank,
Press and Public Relations Division, at the above address, or by fax No. (0 69) 95 66-30 77

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Abstract

Regarding inflation as being a monetary phenomenon in the long-run is a widely-held view in modern macro economics. We analyse this topic by means of a P-star model. Based on the quantity theory of money, this approach explains inflation via a supposed equilibrium price level (P-star), which itself depends on potential output and money. We investigate country-specific models for 110 economies, and also a pooled system thereof. We test for cointegration among money, prices, and real output. Moreover, parameter restrictions for the long-run relationships implied by the monetary theory are tested. Country specific P-star variables are constructed and the cointegration property between prices and the P-star variable is analysed. Along these lines, we find that actual prices and their P-star counterparts are cointegrated at the pooled level and thus demonstrate the importance of money for the development of prices.

Keywords: Quantity theory of money; Panel cointegration analysis; Wild bootstrap inference.

JEL Classification: E31, C22

Zusammenfassung

Die Aussage, dass Inflation langfristig ein monetäres Phänomen ist, ist eine grundlegende Erkenntnis der Makroökonomie. Dieser Sachverhalt wird in dieser Arbeit im Rahmen eines P-Stern-Modells analysiert. Ausgehend von der Quantitätstheorie des Geldes erklärt der P-Stern-Ansatz die Inflationsrate mit Hilfe eines berechneten Gleichgewichtspreisniveaus (P-Stern), das vom Produktionspotential und von einer Geldmenge abhängt. Wir untersuchen länderspezifische Modelle für 110 Volkswirtschaften und ein aus den Gleichungen bestehendes gepooltes System. Es wird auf Kointegration zwischen einer Geldmenge, einem Preisniveau und dem realen Output testet. Weiterhin werden die Parameterrestriktionen, die sich durch die monetäre Theorie ergeben, für die langfristige Beziehung überprüft. Zusätzlich werden nationale P-Stern-Variablen konstruiert und es wird auf Kointegration zwischen diesen und den Preisen getestet. Im empirischen Teil finden wir, wenn die Variablen gepoolt werden, dass die aktuellen Preisniveaus mit den dazugehörenden P-Stern-Variablen kointegriert sind. Dies verdeutlicht die Wichtigkeit des Geldes für die Entwicklung der Preise.

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1 Introduction

It is widely believed in modern macro economics that inflation is a monetary phenomenon. Most central banks monitor monetary aggregates, and major central banks are required to stabilize price movements. For instance, the European Central Bank (ECB) has the primary objective of maintaining price stability in the Euro area (see ECB, 1999). To stress the importance of monetary variables, the ECB gives one monetary aggregate, namely M3, a prominent role in its strategy. An actual growth rate of this aggregate, which exceeds its reference value is taken to indicate that the future inflation rate may be higher than the target level.

The empirical evidence in favour of this inflation explanation is substantiated in two ways. On the one hand, researchers are interested in stable money demand equations for individual countries (e.g. see Lütkepohl & Wolters 1999). Given a stable equilibrium relation among output, prices, interest rates and money, monitoring the latter is sensible when fighting inflation. Moreover, in such a situation it is feasible to control money growth by means of interest rate adjustments. On the other hand, the long-run relationship between money and prices is addressed in the so-called P-star model, which is based on the quantity theory of money. It explains inflation via short-term dynamic factors, and differences between the actual and an equilibrium price level, which itself depends on money and potential output (Hallman et al. 1991, Tödter &

*Helmut Herwartz: Institute of Statistics and Econometrics, Humboldt–University Berlin, Spandauer Str. 1, D - 10178 Berlin, Germany, e-mail: helmut@wiwi.hu-berlin.de. Hans-Eggert Reimers, Hochschule Wismar, University of Technique, Business and Design, Postfach 12 10, D - 23952 Wismar, Germany, e-mail: h.reimers@wi.hs-wismar.de. Part of this research was conducted while the second author was staying at the Deutsche Bundesbank as a visiting researcher. The hospitality of the Bundesbank is greatly appreciated. The paper was presented at seminars of the European Central Bank and of the Economics Institute of the University of Regensburg. We thank the participants, especially Heinz Herrmann and Franz Seitz, for their helpful comments. The views expressed in this study are our own, and not those of the Deutsche Bundesbank.

Reimers 1994, Hoeller & Poret 1991, Kool & Tatom 1994, Gerlach & Svensson 2001).

Compared to a large body of econometric literature on money demand models, only a few empirical studies focus directly on the link between prices and money. Recently, McCandless & Weber (1995), Romer (1996) and Lucas (1996) have graphically illustrated the comovement of money growth and inflation for more than 100 countries. Analysing a large set of macroeconomies the latter contributions are to be recommended for the information base exploited. From an econometric point of view, however, the approach adopted of unconditional empirical moments appears to be inefficient. It is more fruitful to take the dynamic structure of the involved variables into account within a time series modelling framework.

Modern econometric concepts permit joint modelling of short and long-run dynamics, and therefore incorporate the ability of variables to adjust towards long-run equilibrium. Especially if the variables are nonstationary, the concept of cointegration is sensible (Engle & Granger 1987). Furthermore, to improve the power of common test procedures in cointegrated systems it is sensible to enhance the information base by pooling macroeconomic entities. Panel cointegration methods have attracted a lot of interest in the recent econometric literature (Banerjee 1999). Pedroni (1997) proposes a residual based procedure to test for cointegration in pooled dynamic systems. Building on single equation error correction models Herwartz and Neumann (2000) have introduced a bootstrap approach estimating critical values for common tests on cointegration and equilibrium relations at the level of (high dimensional) pooled equations.

Extending the existing literature, we analyse the importance of money in the development of prices where utilization of the panel approach is the main innovation. Firstly, cointegration is tested among money, prices and real output for each country, and at the level of pooled economies. Secondly, restrictions of the cointegrating space are tested for the pooled system. Thirdly, we construct country-specific P-star variables, and, fourthly, test for cointegration between actual prices and the P-star series. Following the P-star approach we find cointegration at the pooled level and thus confirm the predominant role of money in the evolution of prices.

The remainder of the paper is organized as follows. In the next section the long-run relation between money and prices is explained, and testable restrictions are given. In Section 3 the bootstrap method adopted to obtain critical values for particular test statistics is briefly provided. In Section 4 the data and descriptive analyses are given. Section 5 presents empirical results. Section 6 summarizes and concludes.

2 The theoretical model

The long-run relation between money and prices is based on the quantity equation

$$P \times Y = M \times V, \tag{1}$$

where P is the price level, Y is real output, M is the money supply and V is the velocity of money. Owing to the definitions of the variables the relation in (1) is an identity.

By making two simplifying assumptions, the quantity equation becomes a theory of the cause of inflation. Firstly, the velocity of money is regarded as depending on the institutional structure of the financial system that includes the payment system. Since this system might be supposed to be time invariant, it is sensible to treat V as being constant. If, secondly, output is exogenous for money, price changes in money must be reflected in inflation rates. The practical relevance of the latter implication can be illustrated by scatter plots showing average inflation rates against money supply growth (Romer 1996, p. 392, McCandless and Weber 1995, p. 3). Moreover, McCandless and Weber (1995) show that average money growth does not affect average output growth. More recently, McCallum (2001) has emphasized that the unsystematic part of policy instrument variability is small in relation to the variability of the systematic component. Therefore, the investigation of the systematic component should dominate the analysis.

In a growing economy, Y may increase at some steady rate, thereby (partially) "absorbing" money growth. Furthermore, invariance of the velocity of money is a strong assumption, which should be confirmed. As long as output and velocity are in equilibrium, however, equation (1) defines the equilibrium price level

$$P^* = (M/Y^*)V^*, \tag{2}$$

where equilibrium values are indicated by a star (*). P^* aims to measure the price level to be obtained at the actual money holdings if production and velocity are in equilibrium. If P and P^* are nonstationary and cointegrated, and the actual price level is below its equilibrium, a future acceleration of inflation can be expected (Hallman et al. 1991).

The equilibrium price level is not directly observable. To calculate P^*, empirical estimates of potential production and trend velocity are required. In some countries, potential output is estimated, but it is not obvious how to obtain trend velocity. If log velocity in time t ($v_t = \ln V_t$) fluctuates randomly around a constant term, it becomes $v_t = v_0 + \epsilon_t$. If ϵ_t is a stationary zero mean process, the equilibrium level of log velocity is $v^* = v_0$. In some countries, however, the velocities of monetary aggregates have exhibited a marked downward trend in the past. Rather than assuming a deterministic trend, Tödter and Reimers (1994) propose incorporating a stochastic trend of velocity if real money demand is income elastic ($\beta_1 > 1$)

$$m_t - p_t = \beta_0 + \beta_1 y_t + z_t, \tag{3}$$

where y_t is the log of real income (GDP), β_0 is a constant term and β_1 is the long-run income elasticity of money demand. If z_t is a stationary stochastic process with zero mean, equation (3) describes a cointegration relationship. King and Watson (1997) refer to (3) as being a monetary equilibrium condition. Contrary to this long-run relation, a short-run dynamic money demand equation would have to take lagged adjustment as well as interest rates into account.

Combining (1) and (3) yields the following expression for velocity

$$v_t = -\beta_0 + (1 - \beta_1) y_t - z_t. \tag{4}$$

This suggests measuring trend velocity as:

$$\begin{aligned} v_t^* &= -\beta_0 + (1 - \beta_1) y_t^* \\ &= v_0 + (1 - \beta_1) y_t^*. \end{aligned} \tag{5}$$

For $\beta_1 = 1$, this approach encompasses the stationary velocity case. If $\beta_1 > 1$, a declining trend of velocity is induced as long as potential output is growing.

Substituting (5) into the definition of P^* in (2), we end up with the following measure of equilibrium prices:

$$p_t^* = m_t - \beta_1 y_t^* + v_0. \tag{6}$$

Based on equation (15), there are some testable relationships. Firstly, one may be concerned with cointegration between money, prices and real output, since these variables are observable. The cointegrating relationship is written as

$$\beta_2 p_t + \beta_3 m_t + \beta_1 y_t$$

and the first hypothesis we consider is:

$$H_1 : \ p_t, \ m_t \ \text{ and } y_t \text{ are cointegrated.}$$

Secondly, this long-run relation should not affect output, i.e.

$$H_2 : \ \text{output is weakly exogenous.}$$

Thirdly, the coefficients of prices and money should be equal in absolute terms.

$$H_3 : \ |\beta_2| = |\beta_3| \ \text{ price homogeneity or proportionality.}$$

Fourthly, if the cointegration relation is normalized such that the coefficient of the price level is unity, the coefficient of real output is (minus) unity:

$$H_4 : \ |\beta_1| = 1.$$

Condition H_4 implies that the velocity of money is constant in the long run. If $|\beta_1| > 1$, it picks up a decreasing trend in velocity. If an estimate of potential output is available, all these conditions are testable, applying a cointegration test between the equilibrium price level and the actual price level. This test is based on an error correction model for prices involving the price gap. Cointegration is found if the influence of the price gap in the price equation is significant. Therefore, the final hypothesis we investigate is

$$H_5 : \ p_t \ \text{ and } \ p_t^{\star} \ \text{ are cointegrated.}$$

One further remark is make for H_2. If output is weakly exogenous it is assumed that y_t^* is exogenous. The construction of v_t^* via (5) implies that it is likewise exogenous.

3 Methodology

In this paper, we examine the monetary policy hypotheses given in the foregoing section for a panel consisting of 110 economies. For this reason, we adopt a methodological framework recently advocated by Herwartz and Neumann (2000), which is easily implemented for large panels and allows asymptotically efficient inference even in the presence of cross sectional error correlation. The method consists in aggregating OLS-based single equation likelihood ratio (LR-) test statistics over the set of considered economic entities. Both LR-tests on long-run parameters and on weak exogeneity are aggregated without modelling contemporaneous error correlation explicitly. In the presence of cross sectional error correlation, such a pooled LR-statistic is no longer pivotal, and its asymptotic distribution depends on nuisance parameters. Herwartz and Neumann (2000) prove the validity of a bootstrap procedure in estimating critical values for aggregated (dependent) LR-type statistics. As its main advantage the methodology adopted here does not require any first step estimate of the underlying pattern of contemporaneous error correlation. To account explicitly for cross sectional error correlation in the present case considering 110 macroeconomies feasible generalized least squares methods would require the estimation of a covariance matrix of dimension 110×110.

In the following we provide only a brief account of the inference techniques adopted. For a detailed discussion the reader is referred to Herwartz and Neumann (2000). Assuming presample values to be available, we first consider the following single country empirical model for real money (mr_t):

$$\begin{aligned} \Delta mr_t &= \delta_0 + \alpha_1(p_{t-1} + \beta_3 m_{t-1} + \beta_1 y_{t-1}) + \delta_1 \Delta mr_{t-1} \\ &\quad + \delta_2 \Delta p_t + \delta_3 \Delta p_{t-1} + \delta_4 \Delta y_t + \delta_5 \Delta y_{t-1} + u_t. \end{aligned} \tag{7}$$

The adopted single equation approach to estimating long-run equilibrium relations is asymptotically efficient if a set of assumptions can be made (Banerjee et al. 1993, Chapter 6). In order to ensure the white noise property of u_t, the equation in (7) may be conveniently augmented by further (lagged) stationary explanatory variables. The LR-statistics derived from the single equation ECM are typically represented as T times the log ratio of the sum of residual squared errors, estimated via OLS under a particular null hypothesis and its alternative, respectively. With RSS_0 and RSS_1 denoting these estimates the LR-statistic looks formally as follows:

$$\mathrm{LR}_n = T \ln\left(\frac{\mathrm{RSS}_0}{\mathrm{RSS}_1}\right). \tag{8}$$

For some testing issues, it can be shown that LR_n is asymptotically $\chi^2(q)$ distributed, with q denoting the number of excess parameters under the alternative model, compared to the restricted model specified under the null hypothesis.

In Herwartz and Neumann (2000), it is shown that the so-called wild bootstrap (Wu 1986) yields valid critical values if the LR-statistic is obtained from pooling over a set of equations. To be more precise on the issue of pooling, consider now a set of empirical error correction models, for example

$$\begin{aligned}\Delta mr_t^n &= \delta_0^n + \alpha_1^n(p_{t-1}^n + \beta_3^n m_{t-1}^n + \beta_1^n y_{t-1}^n) + \delta_1^n \Delta mr_{t-1}^n \\ &\quad + \delta_2^n \Delta p_t^n + \delta_3^n \Delta p_{t-1}^n + \delta_4^n \Delta y_t^n + +\delta_5^n \Delta y_{t-1}^n + u_t^n,\ n = 1, \ldots, N, \end{aligned} \tag{9}$$

where N denotes the number of equations (countries) in the system. We are interested in testing a specific null hypothesis to hold in the pooled system. Assuming the error terms of the N equations to be contemporaneously uncorrelated, a convenient generalization of the statistic given in (8) is:

$$\mathrm{LR}_N = \sum_{n=1}^{N} \mathrm{LR}_n = T \sum_{n=1}^{N} \ln\left(\frac{\mathrm{RSS}_{0n}}{\mathrm{RSS}_{1n}}\right). \tag{10}$$

In (10), we implicitly assume that T observations are available for each equation. Note that LR_N is easily modified if this assumption is violated. In the case of contemporaneous correlation between equations, critical values of the LR_N-statistic can be obtained by means of a bootstrap procedure. Apart from cross sectional correlation, the wild bootstrap also accounts for heteroskedastic error distributions.

4 Data and descriptive analysis

The monetary policy hypotheses are examined for 110 countries. The data are from the data base of the International Monetary Fund (international financial statistics). To approximate price level and real output, we use the consumer price index (CPI) and real gross domestic product (GDP), respectively. For some countries, the latter variable is not available so that we replace it by nominal GDP deflated by CPI. As the money variable, two alternatives are used. Firstly, we implement the empirical analysis with a narrow definition of money, namely M1. Since notably central banks of industrial countries stress the importance of a broad monetary aggregate, we also used a variable Mb, which is defined as the sum of M1 and so-called quasi money. For some members of the Eurozone, the latter variable cannot be constructed so that we rely on a country-specific broad monetary aggregate (e.g. M3) to be found in the IMF or Deutsche Bundesbank data base.

We investigate annual data for the (maximum) sample period 1960 to 1999. Mostly data are available for (non-overlapping) country-specific subperiods, which cover at least 15 years. All variables are measured in natural logarithms. Given alternative definitions of money we analyse two sets of variables, namely

- S1: M1, CPI and real GDP, and
- S2: Mb, CPI and real GDP.

For Sweden and Italy, data on M1 and Mb are not available, respectively. Thus each set used for the empirical analysis contains 109 countries.

To provide a first look at the data, we show average geometric growth rates of real GDP, consumer prices and the two definitions of money. In the upper panels of Figure 1 scatter plots are given for growth of M1 and Mb versus CPI inflation. To facilitate their interpretation both illustrations also show the 45°-line. Observations for individual countries are close to the diagonal. This result does not depend on the definition of money employed and is confirmed by regressing the average inflation rate on a constant and money growth. Results for such a regression are shown in Table 1. The degree of explanation (R^2) is rather high, and the coefficients of the money variables are close to unity. These results are in line with evidence in de Grauwe and Grimaldi (2001).

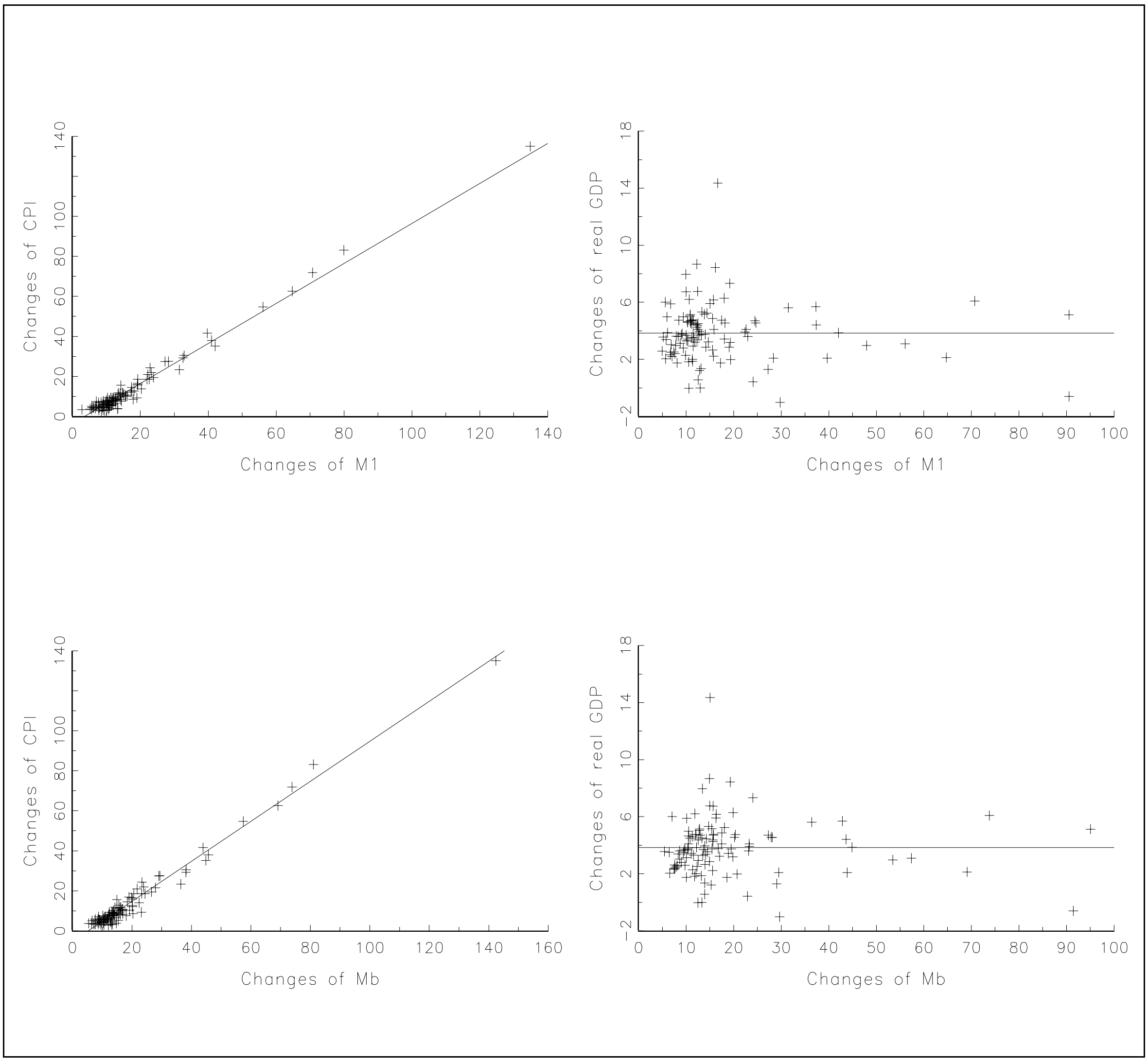

Figure 1: Average yearly growth rates of M1, Mb, consumer price index and real output for a maximal period from 1960 through 1999 for 110 countries.

Considering two subsets, namely the set of countries belonging for more than 10 years to the OECD† and the set of Latin America economies,‡ it is worth emphasizing that there are only small regional differences. Similar charts covering the 1960-90 pe-

†The set OECD includes the countries: Austria, Australia, Belgium, Canada, Denmark, Finland, France, Germany, Greece, Iceland, Italy, Ireland, Japan, Netherlands, New Zealand, Norway, Portugal, Spain, Switzerland, Turkey, United Kingdom, and United States.

‡Latin America comprises Argentina, Bahamas, Bolivia, Brazil, Chile, Colombia, Costa Rica, Dominican Republic, El Salvador, Guatamala, Haiti, Honduras, Jamaica, Mexico, Panama, Paraquay, Peru, Uruguay, Venezuela.

riod can be found in McCandless and Weber (1995). Brazil has the highest inflation rate and money growth rate. Both plots and the regression results seem to confirm a one-to-one relationship between money growth and inflation. Moreover, de Grauwe and Grimaldi (2001) analyse the stability of the money coefficient as a function of the level of inflation. The data are arranged in ascending and descending order of inflation. They show that the money coefficient increases with the average inflation rate. For low inflation countries, the coefficients are not significantly different from zero. Adding high inflation countries to the sample leads to an increasing coefficient. Including all countries, the latter coefficient is close to unity.

Table 1: Regression results for the average inflation rate (real output growth rate) and the growth rates of monetary aggregates.

Vari-ables	Regions								
	All countries			OECD			Latin America		
Δp	C	$\Delta m1$	R^2	C	$\Delta m1$	R^2	C	$\Delta m1$	R^2
	-4.25 (13.25)	1.04 (78.01)	0.983	-2.46 (3.59)	0.94 (17.14)	0.936	-4.21 (6.14)	1.04 (66.97)	0.996
Δp	C	Δmb	R^2	C	Δmb	R^2	C	Δmb	R^2
	-5.29 (13.31)	1.00 (64.95)	0.975	-2.84 (5.92)	0.87 (24.74)	0.968	-5.69 (7.06)	1.00 (58.26)	0.995
Δy	C	$\Delta m1$	R^2	C	$\Delta m1$	R^2	C	$\Delta m1$	R^2
	4.03 (13.69)	-0.011 (0.86)	0.006	3.10 (8.49)	0.038 (1.35)	0.083	3.22 (6.53)	0.013 (0.97)	0.051
Δy	C	Δmb	R^2	C	Δmb	R^2	C	Δmb	R^2
	3.95 (12.86)	-0.006 (0.49)	0.002	3.02 (8.50)	0.036 (1.48)	0.099	3.20 (6.25)	0.012 (0.98)	0.050

Maximum of information period: 1960-1999. Estimated t-values in parentheses. OECD includes countries that have been members of the organization for more than ten years.

The lower panels of Figure 1 show average growth rates of money versus real output for all the countries considered. The overall average growth rate of real output is also displayed. It appears that money growth does not affect the growth of real output. Moreover, output growth is regressed on an intercept and money growth. Regression results are also given in Table 1. Regardless of the monetary aggregate employed its coefficient is not significantly different from zero in a regression covering all countries.

The degree of explanation is small. The latter findings are in line with McCandless and Weber (1995), as well as de Grauwe and Grimaldi (2001). Performing the latter regression for the OECD subset, we find that the coefficient of growth of the broad monetary aggregate is positive and the corresponding t-value and R^2 are higher compared with the remaining cases considered. On the one hand, this result may indicate that, for the industrial countries, money may lose its neutrality. On the other hand, however, the interpretation could be reversed, i.e. the money supply is extended if the growth rate of a country increases.

5 Econometric results

Neglecting the time series dimension of the data, the comparison of average growth rates, as adopted so far, is a weak econometric concept in terms of efficiency. In the following single equation error correction models (ECMs) are used to reexamine long-run links between prices and money as well as money and real output. Its application requires the variables to be integrated of order one and one cointegrating relationship to exist among the variables. These assumptions are tested via the multivariate maximum likelihood methodology (Johansen 1995), which yields evidence of more than one cointegration relation for only 17 countries.[§] For each set of variables, we specify three country-specific ECMs, namely:

$$\Delta mr_t = \nu_1 + \alpha_1(p_{t-1} + \beta_3 m_{t-1} + \beta_1 y_{t-1}) + r_{1t} + u_{1t}, \tag{11}$$

$$\Delta y_t = \nu_2 + \alpha_2(p_{t-1} + \beta_3 m_{t-1} + \beta_1 y_{t-1}) + r_{2t} + u_{2t}, \tag{12}$$

$$\Delta p_t = \nu_3 + \alpha_3(p_{t-1} + \beta_3 m_{t-1} + \beta_1 y_{t-1}) + r_{3t} + u_{3t}. \tag{13}$$

The processes r_{it}, $i = 1, 2, 3$ in (11) to (13) are equation-specific linear combinations of stationary variables (current and lagged values of Δmr_t, Δm_t, Δy_t, Δp_t), the inclusion of which turned out to be necessary so as to yield serially uncorrelated error terms u_{it}. Since the particular composition of r_{it} varies across countries and is informative only for short-run dynamics, we employ this brief notation. Within the empirical exercises, we aim to specify r_{it} as parsimoniously as possible. In fact, each estimated coefficient in $\hat{r}_{it}$ has an absolute $t-$ratio of at least 1.5.

[§]The results are available from the authors upon request.

Table 2: Regression results for the United States

Regres.-sors	$\Delta m1r$	Δmbr	Δy	Δy	Δp	Δp
	Dependent variables					
C	-0.058 (0.44)	0.022 (0.12)	0.163 (1.76)	0.142 (1.62)	-0.149 (2.62)	-0.083 (1.30)
$m1_{t-1}$	-1.651 (7.69)		-1.139 (1.62)		-0.817 (2.84)	
mb_{t-1}		-0.767 (5.98)		6.647 (0.11)		0.327 (1.24)
p_{t-1}	0.286 (4.88)	0.295 (3.76)	0.050 (1.13)	-0.004 (0.10)	-0.093 (3.03)	-0.065 (2.12)
y_{t-1}	1.628 (3.75)	0.330 (0.93)	0.282 (0.19)	16.92 (0.10)	-0.040 (0.07)	-0.686 (1.02)
$\Delta m1r_{t-1}$	0.384 (3.45)					
Δmbr_{t-1}		0.687 (4.39)				
Δp	-1.393 (6.84)					
Δp_{t-1}		0.512 (2.02)	-0.517 (3.70)	-0.470 (4.88)	1.075 (8.92)	0.958 (9.92)
Δy_{t-1}					0.458 (3.79)	0.372 (3.18)
Δy_{t-2}			-0.215 (1.51)			
$\Delta m1_{t-1}$					0.100 (1.79)	
Δmb						-0.114 (1.91)
Δmb_{t-1}				0.191 (2.62)		
R^2	0.71	0.57	0.43	0.53	0.85	0.84
DW	1.97	1.74	2.18	2.21	1.82	1.70

Estimated subset models in the error correction form for the United States. Information period: 1960-1999. Estimated t-values in parentheses.

The two sets of variables differ with respect to the monetary aggregate employed, i.e. the variable m_t is either narrowly ($m_t = m1_t$) or broadly ($m_t = mb_t$) defined. The coefficients $\alpha_i, i = 1, 2, 3$, indicate how the dependent variable responds to lagged violations of the long-run equilibrium relation, and it is expected to be negative (positive) for the first (third) equation. If output is weakly exogenous for inference on the long-run equilibrium, α_2 is expected to be zero (H_2). In some cases, the empirical models are augmented with one or two impulse dummy variables to account for 'extreme' outliers. The overall performance of the specification strategy adopted is monitored by means of the Durbin-Watson (DW) and the Box-Pierce statistic (of order 4 or 8), which we regard as a convenient descriptive statistic for empirical models containing lagged dependent variables. For almost all of the 654 regression models the DW-statistic is close to 2, indicating convenient selections of lagged stationary regressor variables in (11) to (13).

The selected (subset) single equation models are used to implement LR-tests on the level of pooled equations. To provide a particular example for the empirical analysis the results for the US are given in Table 2. The cointegration hypothesis is tested using the t-value of the loading coefficient ($H_1 : \alpha_1 = 0$). In the price equation, it is the coefficient of the lagged price variable. Its t-value is 3.03, which is significant at the 5% level. Weak exogeneity of variables implies that the specified cointegration relationship is not significant in the other equations. Thus the loading coefficient of the cointegration relation should be zero in the output equation ($H_2 : \alpha_2 = 0$). The estimated t-value is 1.13. The hypothesis of weak exogeneity cannot be rejected. For the price homogeneity hypothesis, the coefficient of the money level in price equation should be unity in absolute terms (H_3). Its estimated absolute value is 0.817, which is supposed not to differ significantly from unity. The unrestricted estimate of β_1 is .04: hypothesis H_4 is rejected.

The described tests are conducted for each country. Rejection frequencies for the particular hypotheses are given in Table 3. Three significance levels are distinguished. Alternatively, critical values are taken from the χ^2-distribution and estimated via a bootstrap approach. Given that the former method turned out to be severely oversized for typical sample sizes, it is not surprising that the bootstrap procedure yields rejections of all null hypotheses less frequently. Using critical values from the bootstrap, the hypothesis of no cointegration is rejected for 57 countries at the 5% test level.

Table 3: Frequencies of rejection of cointegration and restriction hypotheses for all countries.

		Asymptotic critical values			Bootstrapped critical values		
Equation	Hypothesis	0.01	0.05	0.10	0.01	0.05	0.10
Set 1 $\Delta m1r$	H_1 : coint.	52	70	77	21	57	68
	$H_3 : \vert\beta_3\vert = 1$	35	46	55	18	36	43
Set 2 Δmbr	H_1: coint.	38	62	71	15	42	61
	$H_3 : \vert\beta_3\vert = 1$	26	38	44	13	26	36
Set 1 Δy	H_2: exog.	19	37	47	7	22	37
Set 2 Δy	H_2: exog.	18	31	42	10	23	33
Set 1 Δp	H_1: coint.	38	56	62	19	43	55
Set 2 Δp	H_1: coint.	32	50	64	11	38	53

Maximum of information period: 1960-1999. Number of countries: 109. Set 1 (2) includes M1, Y and P (Mb, Y and P), respectively. Hypothesis 'coint': null hypothesis of no cointegration is tested in the corresponding equation. The test uses the loading coefficient to check the hypothesis. Hypothesis 'exog': null hypothesis of weakly exogenous variable for the cointegration relationship. The assumed p-values, depending on the asymptotic results (bootstrapped approach), for the critical values are 0.01, 0.05 and 0.10.

Table 4: Trimmed average estimates of the cointegration relationship:
$\alpha_i(p_{t-1} + \beta_3 m_{t-1} + \beta_1 y_{t-1})$

	Set 1			Set 2		
Equation	p_{t-1}	$m1_{t-1}$	y_{t-1}	p_{t-1}	mb_{t-1}	y_{t-1}
Δmr	0.462	-1.067	1.053	0.314	-0.857	0.614
Δy	-0.243	-0.746	0.465	-0.267	-1.120	2.319
Δp	-0.101	-0.679	-0.460	-0.092	-0.763	0.697

Maximum of information period: 1960-1999. Number of countries: 109. Set 1 (2) includes M1, P and Y (Mb, P and Y), respectively.

To give an impression of the coefficient estimates obtained trimmed average values are shown in Table 4. To mitigate the impact of outlying estimates on these averages, the two biggest and two smallest coefficient estimates do not contribute to the reported average values. Note that the long-run parameters are estimated as the ratio of two OLS estimators. If the estimated error correction coefficient is close to zero, this ratio may become very large in absolute value. In general, the values have the expected signs. For example, in the money equations, the average loading coefficient is positive. The coefficient of the money (output) variable in set 1 is -1.067 (1.053), which is close to unity. In Table 5, test results for the pooled system are presented. It is apparent that the hypothesis of no cointegration (H_1) is clearly rejected in both money equations and in the price equations. The hypothesis of price homogeneity, however, is rejected in the money equations (H_3). This is surprising, since the average estimated price coefficient is -1.067 in Set 1 (see Table 4). On the other hand, a few estimates of money demand equations include an inflation rate in the cointegrating relationship (see Lütkepohl & Wolters 1999, or Coenen & Vega 1999). Assuming price homogeneity, the hypothesis of a stationary velocity (H_4) is rejected for both sets (see Table 5). This result corresponds to money demand estimates for some industrial countries where the income elasticity is greater than unity (see for example, Brand & Cassola 2000). Weak exogeneity of output is rejected (H_2). This finding confirms the results of Rapach (1999) rejecting long-run neutrality of money for 14 individual industrialized economies. In contrast, King and Watson (1997) find little evidence against the long-run neutrality of money for the US.

The latter results are also found in the two analysed subsystems (see Table 5). The evidence for the Latin America region is more pronounced in the direction of the results for the entire sample. The hypothesis of a stationary velocity cannot be rejected at the 1% level for Set 1 on the assumption of price homogeneity in the price equation. It is rejected for Set 2. This points in the direction that the velocity of the broadly defined monetary aggregate may decrease in the long run.

Turning to the P-star approach, we estimate equilibrium prices which rely on estimates of potential output. It is determined by using an extended exponential smoothing filter (see Appendix B). Tödter (2000a) shows that this filter is symmetric. The original and the filtered series are equal on average. On the hypothesis of price homogeneity or proportionality, the estimated income elasticities of the money equations are used to calculate the P*-series (ECM-approach). Alternatively, the income elasticities are de-

Table 5: Test results of the pooled ECM-equations.

Region	Equation	Hypothesis	Set 1 test value	Set 1 p-value	Set 2 test value	Set 2 p-value
World	Δmr	H_1 : coint.	581.40	0.000	420.48	0.000
		$H_3 : \|\beta_3\| = 1$	1033.36	0.000	857.55	0.000
	Δy	H_2 : exog.	497.61	0.000	509.07	0.000
	Δp	H_1 : coint.	745.68	0.000	680.10	0.000
		$H_3 : \|\beta_3\| = 1$	1296.58	0.000	1195.39	0.000
		$H_4 : \|\beta_1\| = 1\|\|\beta_3\| = 1$	751.46	0.000	564.67	0.000
Latin America	Δmr	H_1 : coint.	135.88	0.000	111.38	0.000
		$H_3 : \|\beta_3\| = 1$	193.35	0.000	189.92	0.000
	Δy	H_2 : exog.	91.68	0.000	67.97	0.002
	Δp	H_1 : coint.	136.52	0.000	91.68	0.000
		$H_3 : \|\beta_3\| = 1$	188.18	0.000	163.77	0.000
		$H_4 : \|\beta_1\| = 1\|\|\beta_3\| = 1$	192.46	0.000	141.21	0.000
OECD	Δmr	H_1 : coint.	143.41	0.000	84.05	0.000
		$H_3 : \|\beta_3\| = 1$	277.21	0.000	134.90	0.000
	Δy	H_2 : exog.	64.60	0.002	63.94	0.000
	Δp	H_1 : coint.	169.23	0.000	141.03	0.000
		$H_3 : \|\beta_3\| = 1$	267.91	0.000	262.84	0.000
		$H_4 : \|\beta_1\| = 1\|\|\beta_3\| = 1$	52.16	0.017	60.41	0.001

Maximum of information period: 1960-1999. Number of countries: 109. Set 1 (2) includes M1, P and Y (Mb, P and Y), respectively. The p-values are calculated by the bootstrap approach.

termined by the static Engle-Granger cointegration regression (Engle and Granger 1987):

$$m_t - p_t = c_0 + \beta_1 y_t + z_t, \tag{14}$$

where z_t is a stationary process. The estimated coefficient $\hat{\beta}_1$ is used to calculate the P*-series. The specification of the dynamic price equation is:

$$\Delta p_t = \delta_0 + \alpha_4(p_{t-1} - p^*_{t-1}) + \delta_1 \Delta p_{t-1} + \delta_2 \Delta p_{t-2} + \delta_3 \Delta p^*_t + u_{4t}. \tag{15}$$

where u_{4t} is a white noise process. Equation (15) includes the price gap $(p_{t-1} - p^*_{t-1})$ and dynamic factors. In contrast to Gerlach and Svensson (2001), who use quarterly observations on Δp^*_{t-1}, we employ annual data. The model (15) allows us to test the hypothesis:

$$H_5 : \alpha_4 = 0 .$$

Figure 2 gives the Box-plots of estimated income coefficient ($\hat{\beta}_1$). It is seen that the ECM-approach yields more extreme values of $\hat{\beta}_1$ compared with the static regression (14).

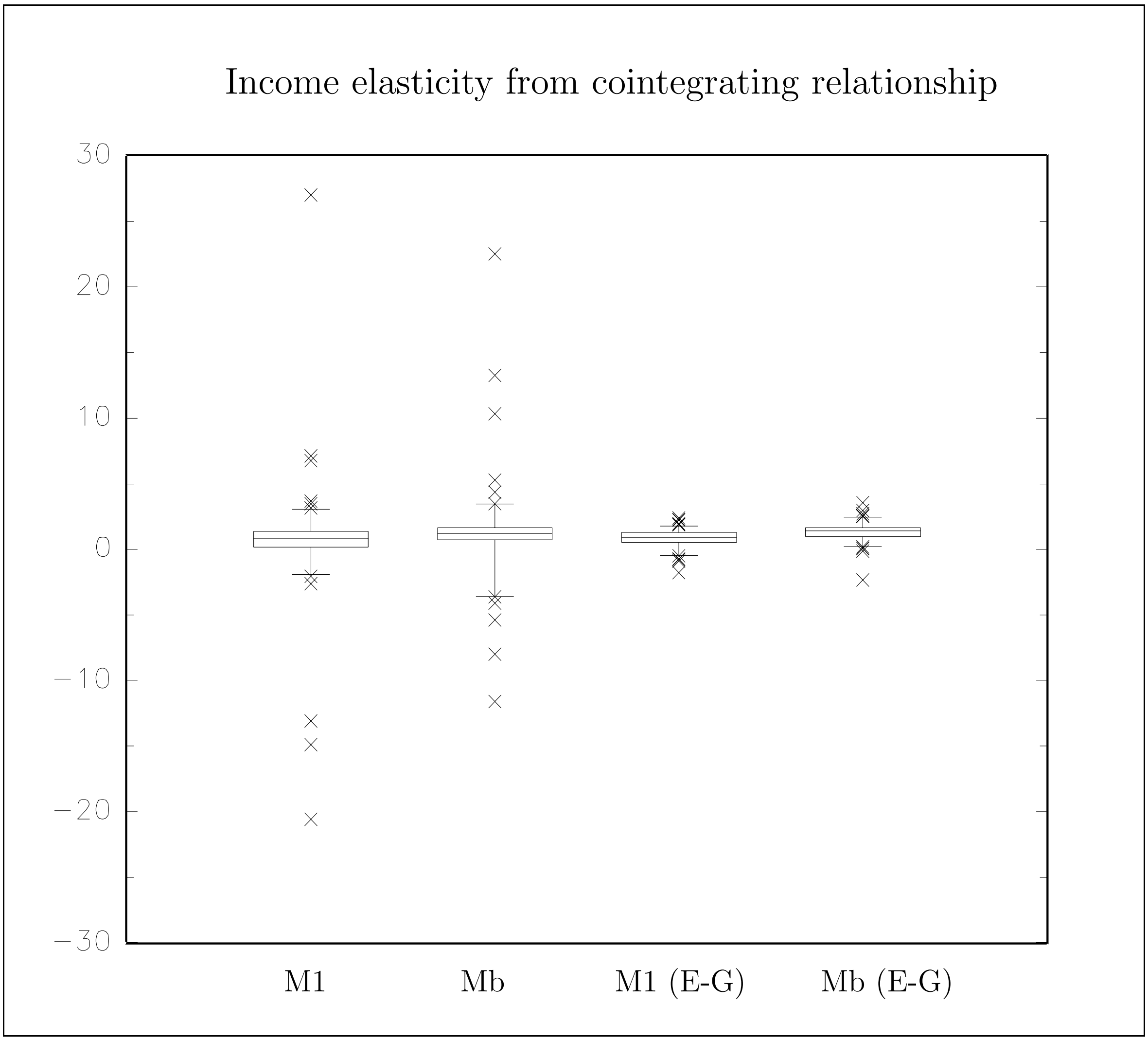

Figure 2: Box-Plots of estimated income elasticity β_1 of the restricted cointegration relationship $(m + p + \beta_1 y)$ or of Engle-Granger cointegration relation. Upper bound (lower bound) corresponds to the 95th (5th) quantil.

Table 6: Test results of H_5 for the pooled P* equations.

Region	Regression	Set 1 $\tilde{\alpha}$	Set 1 test value	Set 1 p-value	Set 2 $\tilde{\alpha}$	Set 2 test value	Set 2 p-value
World	ECM	-0.123	596.05	0.00	-0.119	496.02	0.00
	E-G	-0.163	700.31	0.00	-0.174	638.27	0.00
Latin	ECM	-0.195	116.81	0.00	-0.183	130.99	0.00
Amer.	E-G	-0.217	137.56	0.00	-0.208	126.13	0.00
OECD	ECM	-0.040	103.10	0.00	-0.051	48.71	0.03
	E-G	-0.071	99.88	0.00	-0.123	94.35	0.00

Maximum of information period: 1960-1999. Number of countries: 109. $\tilde{\alpha}$: average value of $\hat{\alpha}_i$ that is corrected by outliers. Regression: Indication how β_1 is estimated, ECM: ECM-approach, E-G: Engle-Granger approach. Set 1 (2) includes M1 (Mb), respectively. The p-values are calculated by the bootstrap approach.

Table 6 presents the results for the level of pooled economies. It is apparent that the hypothesis of no cointegration is clearly rejected in price equations for the largest system. The price gap is significant at all reasonable test levels. Its influence on the development of prices has the expected (negative) sign. The differences with respect to the alternative monetary aggregates are small. The test statistics are higher for the equation based on the Engle-Granger approach than for the ECM approach. Accordingly, the mean adjustment coefficient is smaller, using estimates of β_1 from (14). It indicates that more than 15 percent of deviations from the equilibrium value vanish in one year. This value is even higher in the Latin American subsystem. The adjustment speed is lower for the OECD countries. Following the ECM-approach (Engle-Granger regression) we find that only 5 (12) percent of a lagged disequilibrium vanishes in one year.

Following de Grauwe and Grimaldi (2001) we examine the stability of the P-star relation as a function of the level of inflation. For this purpose, individual countries are arranged according to their average inflation rates. The number of pooled equations is increased stepwise. To implement the empirical model, we use income elasticities estimated via the static regression (14). Figure 3 displays the estimates of the average adjustment coefficient and the p-value for the hypothesis of no-cointegration obtained

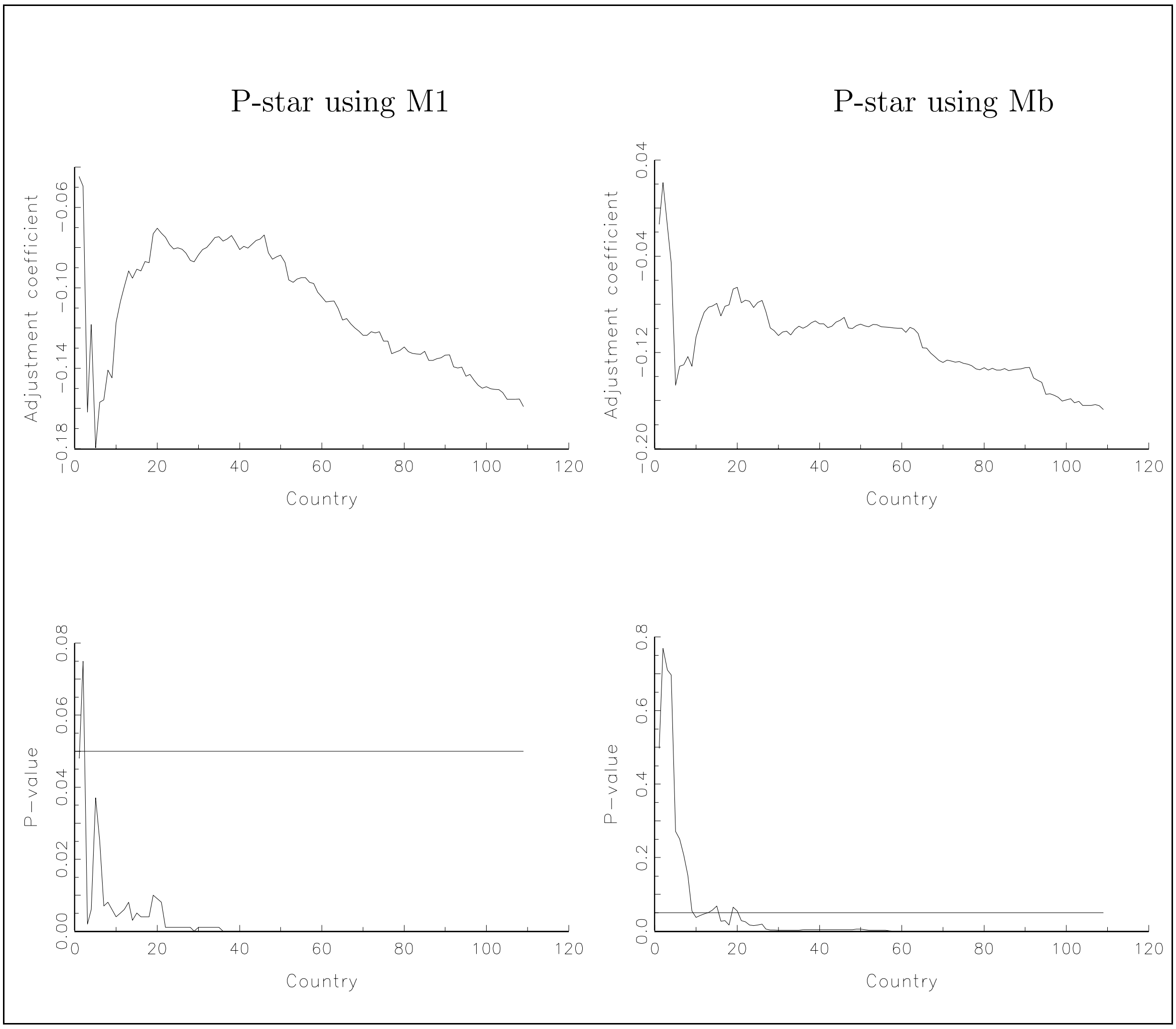

Figure 3: Recursive estimates of the adjustment coefficient and its p-value in the inflation equation using M1 (Mb) where the countries arranged according to their average inflation rate.

from the bootstrap. Apparently, the absolute adjustment coefficient increases if countries with higher inflation are added to the sample. It takes more time in low inflation countries to obtain the long-run relation than in high inflation countries. If the equilibrium price level constructed with Mb is considered, we find that the cointegrating relationship is not significant in countries having very low inflation rates (see the lower right-hand panel of Figure 3). This finding changes if the equilibrium price level calculated by M1 is considered where the cointegrating relationship is also significant in low inflation countries (see the lower left-hand panel of Figure 3). This result indicates that the proportionality between money and prices also appears under these economic conditions. Note that proportionality is reflected by variables measured in levels. By merely relying on (average) growth rates, such a relationship cannot be uncovered.

6 Conclusion

Using the P-star model framework, which bases on the quantity theory of money and allows us to estimate a nonstationary velocity of money, we analyse the development of prices in 110 macroeconomies. Using annual data, the observation period is, at most, 1960 through 1999. To improve the power of the tests, a panel cointegration approach is used. From a methodological point of view, the analysis permits us to consider economic relationships for variables in levels, rather than relying on descriptive comparisons of average growth rates. Critical values for LR-tests are determined by means of a bootstrap procedure.

At the pooled level, we find evidence of cointegration relating actual prices and an equilibrium price level termed P-star. It is found for the entire sample, as well as for two subsets, the OECD and Latin America. The long-run link is obtained for high inflation countries and, in general, for low inflation countries. With respect to monetary policy, our results underscore that it is necessary for central banks to monitor the development of monetary aggregates. Monetary aggregates determine the price level in the long run, and the control of its development is a cornerstone for achieving price stability in the long run.

Appendix A: The countries analysed:

Algeria, Argentina, Australia, Austria, Bahamas, Bahrain, Bangladesh, Belgium, Belize, Bhutan, Bolivia, Botswana, Brazil, Burkina Faso, Burundi, Cameroon, Canada, Central African Republic, Chad, Chile, Colombia, Congo (Democratic Republic of), Congo (Republic of), Costa Rica, Cote d Ivoire, Cyprus, Denmark, Dominica, Dominican Republic, Ecuador, Egypt, El Salvador, Ethiopia, Fiji, Finland, France, Gabon, Gambia, Germany, Greece, Guatemala, Haiti, Honduras, Hungary, Iceland, India, Indonesia, Iran, Ireland, Israel, Italy, Jamaica, Japan, Jordan, Kenya, Korea, Kuwait, Lesotho, Madagascar, Malawi, Malaysia, Mauritania, Mexico, Morocco, Nepal, Netherlands, New Zealand, Nicaragua, Niger, Nigeria, Norway, Pakistan, Panama, Paraguay, Peru, Philippines, Portugal, Qatar, Rwanda, Samoa, Saudi Arabia, Senegal, Seychelles, Sierra Leone, Singapore, South Africa, Spain, Sri Lanka, Sudan, Swaziland, Sweden, Switzerland, Syria, St. Kitts-Nevis, St. Lucia, St. Vincent, Tanzania, Thailand, Togo, Tonga, Trinidad, Tunisia, Turkey, United Arab Emirates, United Kingdom, United States, Uruguay, Venezuela, Zambia, Zimbabwe.

Appendix B: Estimation of Potential Output

The estimation of potential output Y^* is conducted by a statistical procedure. Following Tödter (2000a), the procedure is derived from the function:

$$Z := \underset{(\hat{y}_t, c_1)}{\text{Min}} \left(\frac{\lambda}{2} \sum_{t=2}^{T} (\hat{y}_t - \hat{y}_{t-1} - c_1)^2 + \frac{1-\lambda}{2} \sum_{t=1}^{T} (\hat{y}_t - y_t)^2 \right)$$

The first term reflects the smoothness of the filtered series, and the second term gives the adjustment of the estimated series on the observed series. The first-order conditions are determined by differencing the function to all $\hat{y}_t$ and c_1. The conditions imply that the intercept term c_1 may be determined by the following nonparametric estimate:

$$\hat{c}_1 = \frac{1}{T-1} \sum_{t=2}^{T} (\hat{y}_t - \hat{y}_{t-1}) = \frac{\hat{y}_T - \hat{y}_1}{T-1}$$

The filtered series is:

$$Y^* = A^{-1}Y$$

where

$$A = \frac{1}{1-\lambda}\begin{pmatrix} 1-\frac{\lambda}{T-1} & -\lambda & 0 & 0 & & \cdots & \frac{\lambda}{T-1} \\ -\lambda & 1+\lambda & -\lambda & 0 & & \cdots & 0 \\ 0 & -\lambda & 1+\lambda & -\lambda & & \cdots & 0 \\ \vdots & & & \ddots & \ddots & & \vdots \\ 0 & & & & 1+\lambda & -\lambda & 0 \\ 0 & & & & -\lambda & 1+\lambda & -\lambda \\ \frac{\lambda}{T-1} & 0 & & \cdots & 0 & -\lambda & 1-\frac{\lambda}{T-1} \end{pmatrix}$$

The filter, which is denoted as extended exponential smoothing, is:

$$Y_t^* = \frac{2\lambda}{1+\lambda}\left(\frac{Y_{t-1}^* + Y_{t+1}^*}{2}\right) + \frac{1-\lambda}{1+\lambda}Y_t$$

for $t = 1, \cdots, T-1$ and $\lambda = \frac{1+\theta}{2}$, where θ is the first-order autocorrelation coefficient of the variable w_t

$$Y_t = Y_{t-1} + c_1 + w_t$$

The series w_t is the OLS-residuals of this equation. Hence, this implies another approach to estimate c_1 (see Tödter, 2000b). This equation allows us to account for impulse dummies at a known break point. If a series includes an impulse dummy, the filtering step is adjusted. First, the series is corrected by the impulse dummy. Second, the filtered series is determined. Third, the effect of the impulse dummy is added to the filtered series.¶

¶The countries including an impulse dummy are Argentina, Belgium, Gabon, Germany, Lesotho, Netherlands, Samoa, Sri Lanka and Tanzania.

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