# *Macroeconomic Rationality and Lucas's Misperceptions Model: Further Evidence from Forty-One Countries*

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**Abstract**. Several researchers have examined Lucas's misperceptions model as well as various propositions derived from it within a cross-section empirical framework. The cross-section approach imposes a single monetary policy regime for the entire period. Our paper innovates on existing tests of those rational expectations propositions by allowing the simultaneous effect of monetary and short run aggregate supply (oil price) shocks on output behavior and the employment of advanced panel econometric techniques. Our empirical findings, for a sample of 41 countries over 1949 to 1999, provide evidence in favor of the majority of rational expectations propositions.

*JEL Classification*: E12; E13

*Key Words*: Lucas's misperception model; rational expectations propositions; real shocks; panel data

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## Introduction

Students of macroeconomics know that the rational expectations revolution, beginning in the late 1960s and early 1970s, quickly replaced the modeling of expectations as an adaptive process. Today, both new classical and new Keynesian theorists adopt rational expectations as a fundamental premise. From the new classical side of the methodological debate, Lucas's misperceptions model (Lucas, 1972, 1973, 1975) generates one of the core building blocks - the Lucas supply curve - that is now taken for granted by both sides.[1]

The concept of rational expectations, much like motherhood, commands support from economists of all persuasions based on the soundness of its theoretical justification. In fact, the rational expectations hypothesis was so intuitively appealing that it generated strong adherence to the premise based solely on its methodological arguments. Nonetheless, good methodology requires good empirical testing. An important, albeit no longer active, literature exists that tests Lucas's misperceptions model and its implications. Our paper adds to this important literature.

The key empirical implication of the misperceptions model implies that the slope of the Lucas supply curve (i.e., the trade-off between nominal disturbances and real economic activity) depends on the variability (i.e., uncertainty) of nominal disturbances. Using cross-section data, Lucas (1973), Barro (1976), Alberro (1981), Attfield and Duck (1983), and Hercowitz (1983) find a significant negative relationship for that hypothesis.[2] In particular, by regressing detrended real income on changes in nominal income, which proxies for nominal disturbances, those authors find that the trade-off coefficient negatively correlates with the variance of the inflation rate. According to Lucas (1973), such evidence should cause the rejection of macromodels in which the reduced-form trade-off between nominal shocks and real output remains stable with respect to the parameters of the process governing these shocks.[3]

Kormendi and Meguire (1984) (hereafter simply K-M) test Lucas's proposition using cross-section data. They reject standard reduced-form macromodels in favor of rational expectations macromodels. Their methodology suffers, however, from a major deficiency. To wit, they assume that the same money supply process holds for the whole sample, a not easily defended assumption.

Froyen and Waud (1980) provide the first intracountry analysis using an intertemporal approach. Their results find a negative association between the inflation variance and the trade-off coefficient between real GNP and the change in nominal GNP over 1957 to 1976. Their findings, however, do not accommodate important supply shocks that occurred during the 1970s. More recently, Froyen and Waud (1984) correct that problem and examine the output-inflation trade-off that incorporates demand variability shocks, supply shocks (i.e., proxied by oil price shocks), and inflation volatility. Their new results continue to support Lucas's hypotheses even accounting for oil price shocks.

Katsimbris (1990a, 1990b) uses intracountry time-series data from 1954 to 1985 to investigate the relationship between the output-inflation trade-off and the variances of the rate of change of nominal income and the price level as well as the relationship between the variance of the inflation rate and the variance of the rate of change of nominal income. His results, in contrast, provide little support for those rational expectations relationships.

Finally, Poirer (1991) investigates the effect of anticipated as well as unanticipated money changes on real output, employing a Bayesian-predictivist approach and permitting alternative monetary regimes. His results for the 47-country sample support neutrality of both anticipated and, surprisingly, unanticipated money.

Our paper provides additional evidence on the relationships examined by K-M (1984) with a sample of 41 countries and an extended time period. The paper also employs a more realistic money supply process - which is one of the main contributions of this study - and avoids misspecification problems by taking explicitly into consideration the effect of short-run aggregate supply disturbances on real output. Finally, the paper performs a pooled (panel) data methodology by using recent advanced econometric techniques. Section 1 describes the rational expectations propositions. Section 2 outlines the employed methodological issues and provides evidence on testing the rational expectations propositions. Finally, section 3 concludes.

## 1. Lucas's Rational Expectations Proposition

Lucas's (1973) aggregate supply model is given as follows:

$$Y_{ct} = a_0 + a_1\, \Delta X_t + a_2\, Y_{c,t-1} + u_t \qquad (1)$$

where $Y_{ct}$ is detrended real output, $\Delta X_t$ is the change in nominal income (a proxy for nominal disturbances), and $u_t$ is the random error. The parameter $a_1$ ($0 \leq a_1 \leq 1$) measures the trade-off between real output and changes in nominal income, while $a_2$ ($|a_2| < 1$) measures the speed of adjustment. This output-inflation trade-off reflects the suppliers' aggregate-relative confusion. That is, with incomplete information, individuals confuse aggregate and relative price movements, resulting in a non-vertical short-run Phillips curve.[4]

Higher volatility in nominal aggregate demand reduces the observed output-inflation trade-off, known as the Lucas variability hypothesis. That hypothesis implies that there exists a negative correlation between the output-inflation trade-off ($a_1$) and the variance of the rate of change of nominal income $\sigma_{\Delta x}^2$. Lucas (1973) tests this hypothesis using data from 18 countries over 1953 to 1967 and finds that countries with the highest $\sigma_{\Delta x}^2$ exhibit a much smaller output-inflation trade-off.

Alberro (1981) extends Lucas's empirical analysis by enlarging the sample to include 49 countries. His results sustain the Lucas hypothesis for 1953 to 1969. Abrams *et al*. (1983), using a different assumption about the trade-off parameter, also support the Lucas hypothesis. Jung (1985) conducts an analysis for 56 countries that confirms the Lucas hypothesis, albeit weaker confirmation for less-developed than developed countries is found.

Lucas's methodology, as noted by K-M (1984), exhibits the following problem: the negative relationship between $a_1$ and $\sigma_{\Delta x}^2$ holds even absent rational expectations. Moreover, the tests of the Lucas's variance hypothesis do not allow variation of policy through time within a given country. K-M (1984) avoid those problems by using a sample of 47 countries to allow different policy regimes. In addition, they combine Barro's (1977, 1978) method of estimating the effect of unanticipated money supply changes on real output with Lucas's approach "of drawing inferences from a cross-section of policy regimes". To this end, K-M employ a variant of the simple two-equation model.

## 2. The Data and Empirical Analyses

*The Data*

We use annual data on real output (Y) measured by GDP at constant 1990 prices, nominal income measured by GDP at current prices, money supply (*M*) defined as M1, the price level (*P*) defined as the consumer price index, and the oil price (*OP*) deflated by a general price index.

The data come from the *International Financial Statistics* CD-ROM, which incorporates all revisions to the historical data and cover 1949 to 1999. The included countries possess continuous annual series of real output, the money supply, and commodity prices. The country group employed contains virtually the same countries used by K-M (1984). The Appendix provides details. Throughout the paper, lower case letters indicate variables, but not parameters, measured in natural logarithms. We employ the MicroFit software in the empirical analysis. Finally, we express our gratitude to Professor Pedroni (Indiana University) for providing the software that performs the panel analysis.

*Time Series Integration Analysis*

Before proceeding to the main empirical analysis, we perform time-series, unit root tests, developed by Perron (1988), with and without a time trend ($\tau_\tau$ and $\tau_\mu$) for the *m*, *y*, and *op* variables for each country over 1949 to 1973, 1949 to 1974, ..., and 1949 to 1999. In all cases, we do not reject the unit-root hypothesis in levels at the 5-percent level. Using first differences, we reject the unit-root hypothesis in all cases.[5]

*Estimating Monetary Regimes*

Empirical results depend crucially on model specification. Sensitivity analysis over various ranges of different plausible specifications generates more information on the robustness of certain empirical relationships (Leamer, 1982). For our purposes, we use a bivariate Vector Error Correction Vector Autoregressive (VECVAR) model to generate recursively monetary regimes, first estimated over 1949-1973, next, over 1949-1974, and so on.[6] In particular, the empirical model is a VECVAR model with two variables, namely, real income and money. The reasons for using a VECVAR is first to capture the endogeneity of money and second that non-reported preliminary empirical findings provided support to the presence of a cointegrating relationship between output and money in all countries under investigation.

The first step of the empirical analysis implements the basic VECVAR model (Karras, 1996):

$$\Delta x_t = \Pi_1 \Delta x_{t-i} + \Phi_1 D_t + \Phi_2 EC_{t-1} + rx_t \qquad (2)$$

where $x_t$ is a 2x1 vector of real income and money, $D_t$ is a set of dummy variables capturing changes in the monetary regime (defined below), *EC* is the error correction term (the residuals from the cointegrating vector), $\Pi_1$, $\Phi_1$, $\Phi_2$ are parameters to be estimated, and $rx_t$ is a vector of random errors.

We allow changes in monetary regimes to assess the effect of changes in the operating targets of the monetary authorities on the stability of the relations under investigation. The practice of the monetary policy implementation exhibits two types (at least for the developed countries) of monetary regimes: money supply control regimes and interest rate control regimes. If the central bank controls the money supply, then it adds high-powered money in order to achieve a target for its money growth rate. By contrast, if the central bank controls interest rates, then the monetary response probably differs substantially.

By permitting changes in the monetary regime, we also allow changes in how central banks respond to observed variables that determine the demand for money. That is, the ECVAR model determines country specific monetary shocks jointly with the variables that specify the course of money demand (i.e. output).[7]

Therefore, we evaluate Lucas's propositions while controlling for such changes on a country-by-country basis. In particular, we generate a series of unexpected country-specific monetary disturbances rm (i.e., the residuals from the VECVAR equation that describes the evolution of $\Delta m$) and their variance $\sigma^2_{rm}$.[8] Also, note that CUSUM and CUSUMSQ tests indicate no structural instability.[9]

Next, following K-M's method, we adopt a variant of Barro's real income equation as follows:

$$\Delta y_t = b_0 + \sum_{i=0}^{2} b_{i1}\, rm_{t-i} + \sum_{i=1}^{q_1} b_{i2} \Delta y_{t-i} + rr_t \qquad (3)$$

where rm are the residuals from the VECVAR equation that describes the evolution of Δm and rr are the residuals in equation (3). We estimate equation (3), which, except for notation, equals equation (8) in K-M (1984), using OLS. But, we also allow for different lag lengths $q_1$; where we determine $q_1$ through the Final Prediction Error (FPE) or Akaike information criterion (AIC).

Our study uses Barro's two-step estimator method, because the Monte Carlo results by Hoffman *et al.* (1984) and Hoffman (1987) support the consistency of the two-step OLS estimates out of the output equation (K-M, 1991). As in K-M (1984, 1991), we also permit the "peak effect" of a monetary shock on real output to occur within at most two years after the shock. That is, define $x$ as the sum of the first two coefficients of monetary shocks (i.e., $x = b_{01} + b_{11}$). We adopt this assumption for all countries in the sample only because in the majority of cases, the empirical results support such a lag pattern. Then we test whether the "peak effect" of a monetary shock on real output (i.e., $x$) correlates negatively with the variability of the monetary disturbances ($\sigma^2_{rm}$). In other words, we test the strong form of the rational expectation hypothesis (a negative relationship between $x$ and $\sigma^2_{rm}$).

Barro (1976, 1980) argues that monetary volatility makes efforts to identify relative price signals more difficult. Absent that negative relationship, we reject the rational expectations hypothesis. In addition, Lucas (1973) and Barro (1976) suggest a non-linear relationship. Therefore, for our empirical purposes and following K-M (1984), we employ a semilog version of the relevant relationships. In addition, we test the validity of other rational expectations propositions through the bivariate relationships, if any, between the logarithm of $x$ *(lnx)* and the following: the variance of unexpected money shocks ($\sigma^2_{rm}$), the mean of actual money growth ($\mu_{\Delta m}$), and the variance of actual money growth ($\sigma^2_{\Delta m}$). The relationship between *lnx* and the mean of money growth captures the rational expectations hypothesis that expected monetary policy exhibits no effect on real output (i.e., the neutrality hypothesis). Accordingly, monetary neutrality implies that real output growth does not correlate with the expected growth rate of the money supply (Barro, 1976; 1980).

*An Alternative Real Income Equation*

No reason exists, however, to believe that the aggregate supplies in individual countries remained stable over the sample period, especially during the 1970s. Hall (1990) finds that oil prices exert a negative effect on total factor productivity, while Dotsey and Reid (1992) emphasize the adverse effect of oil price shocks on real output fluctuations. Those arguments provide a legitimate reason to let oil price shocks proxy real shocks, at least for less-developed countries where competitive markets do not exist.

Given those arguments, we repeat our empirical analysis under the hypothesis that crude oil price shocks affect real economic activity from supply-side disturbances. Thus, equation (3) expands to become:

$$\Delta y_t = b_0 + \sum^{2} b_{i1}\, rm_{t-i} + \sum^{q_2} b_{i2} \Delta y_{t-i} + \sum^{q_3} bi_3\, \Delta op_{t-i} + rr_{1t} \qquad (4)$$

$i=0 \qquad i=1 \qquad i=0$

where $\Delta op$, is the excess of the growth rate of the world price of oil (which does not include the cost of transporting oil or refined products to the nation under study) over the nation's rate of inflation. We include oil prices in equation (4) as a proxy for real (supply) disturbances. Finally, $rr_{1t}$ are the residuals in equation (4) (i.e. real output disturbances). Once again, we determine the optimal number of lags for $q_2$ and $q_3$ through the AIC. We also introduce a dummy variable that defines a break at 1980 when OPEC lost its ability to maintain stable oil prices.

*Panel Integration*

We test the null hypothesis of non-stationarity versus the stationary alternative, using panel data sets for the total sample of 41 countries as well as for two subgroups of developed countries (DCs) and less-developed countries (LDCs). We employ the group-mean panel unit root-test (or 't-bar' test) of Im, Pesaran and Shin (1995). This test calculates the Augmented Dickey-Fuller (ADF) statistic for each country (Dickey and Fuller, 1981) and allows each member of the cross section to possess a different autoregressive root and different autocorrelation structures under the alternative hypothesis. Table 1 reports the results without and with a trend. We cannot reject the hypothesis that variables $m$, $y$, $op$, $lnx$, $\sigma^2_{rm}$, $\mu_{\Delta m}$, and $\sigma^2_{\Delta m}$ contain a unit root at the 1-percent significance level in all three sub-cases: all countries, only developed countries, and only less developed countries. Employing first differences, we reject unit-root non-stationarity at the 1-percent significance level, suggesting that the variables $m$, $y$, $op$, $lnx$, $\sigma^2_{rm}$, $\mu_{\Delta m}$, and $\sigma^2_{\Delta m}$ are I(1) variables in all three cases. Finally, note that the sequence of moments $\sigma^2_{rm}$ and $\sigma^2_{\Delta m}$ behave as a unit root process. That is, monetary regime shifts generate heteroskedastic data series and produce non-stationary behavior of those variabilities.

*Hypotheses Testing (Panel Cointegration-All countries)*

*Lucas's Variance Hypothesis: Long-Run Negative Relationship between lnx and* $\sigma^2_{rm}$

First, we calculate $\sigma^2_{rm}$ as the variance of the residuals from the VECVAR monetary regimes over 1949 to 1973, 1949 to 1974, ..., and 1949 to 1999. Then, we calculate the natural logarithm of the "peak effect" (*lnx)* from equation (3) over 1949 to 1973, 1949 to 1974, ..., and 1949 to 1999, by taking the sum of the first two coefficients of monetary shocks in equation (3). Once we establish the order of panel stationarity, we can move to a panel cointegration approach. The panel cointegration technique makes use of a residual-based ADF test (Pedroni, 1995; 1997).

This technique employs the residuals of the long-run model (for a panel of N countries and T time observations).

The estimated cointegrating relationship is as follows:

$$lnx_{it} = \beta_{1i}\, \sigma^2_{rmit} + \varepsilon_{it} \qquad (5)$$

where $i$ runs from 1 to N countries and t runs from 1 to T observations. This relationship estimates the long-run relationship. The term $\varepsilon_{it}$ estimates the deviation from the modeled long-run relationship.

If the series are cointegrated, this term is a stationary variable. In other words, we establish stationarity by showing that $\rho_i < 1$ in:

$$\varepsilon_{it} = \rho_i\, \varepsilon_{it-1} + \xi_{it} \qquad (6)$$

The null hypothesis, associated with Pedroni's test procedure, is that $\rho_i = 1$. In other words, the null hypothesis is equivalent to testing the null of nonstationarity (no cointegration) for all i.

Pedroni (1999) provides a total of seven tests of the null of no cointegration, of which four involve pooling on the within dimension (panel tests) and three on the between dimension (group mean tests). Both tests include non-parametric statistics analogous to the Phillips and Perron rho statistic, as well as a parametric t-statistic analogous to the Augmented Dickey-Fuller t-statistic.

The cointegration results appear in Table 2. The results confirm that the panel (without and with oil shocks) is stationary. In other words, the results indicate that $lnx$ and $\sigma^2_{rm}$ share a long-run cointegrating relationship in both cases. Given cointegration, we estimate the long-run relationship through the Dynamic OLS (DOLS) approach due to Stock and Watson (1993). This approach regresses an I(1) variable on other I(1) variables plus lags and leads of the first-differences of the I(1) variables. The inclusion of the first-differenced variables eliminates any possible bias resulting from correlation between the error term and the I(1) variables. We also calculate corresponding robust standard errors through an adjustment suggested by Newey and West (1987), with Bartlett weights and a truncation lag of 4. We employ the DOLS regression by adding one lag and one lead of the first difference of the right-hand side variable to the equation.

**Without oil shocks**

| | |
|---|---|
| Coefficient: | - 0.163 |
| T-statistic | (-5.05)* |

$R^2$ 0.559

$F\sigma^2_{rm}$ = 77.09[0.0]

$F\sigma^2_{rm}$ = 1.52[0.18]

where the first F-test indicates that the coefficients are jointly significant across countries, while the second F-test indicates that the coefficients are equal across countries, indicating homogeneity in the cointegrating vector. In other words, evidence exists to support one common cointegrating vector among the variables in the panel. Figures in brackets indicate p-values.

**With oil shocks**

Coefficient - 0.241

T-statistic (-5.83)*

$R^2$ 0.651

$F\sigma^2_{rm}$ = 45.92[0.00]

$F\sigma^2_{rm}$ = 0.73[0.51]

where the numbers in brackets indicate p-values. The empirical findings provide support for the negative correlation between ln$x$ and $\sigma^2_{rm}$ for the entire sample in both cases.

*The Neutrality of Anticipated Money Growth: No Long-Run Correlation between lnx and $\mu_{\Delta m}$*

The panel cointegration findings (Table 2) confirm that the panel (without and with oil shocks) is nonstationary. In other words, the results indicate that *lnx and* $\mu_{\Delta m}$ do not share a long-run cointegrating relationship in both cases.

*The Neutrality of the Variance of Anticipated Money Growth: No Long-Run Correlation between lnx and $\sigma^2_{\Delta m}$*

Panel cointegration findings (Table 2) indicate that *lnx and* $\sigma^2_{\Delta m}$ do share a long-run cointegrating relationship in both cases.

*Testing the Robustness of the Results*

Next, we test the above panel rationality hypotheses for two subgroups, one that involves only developed countries (DCs) and one that involves only less developed countries (LDCs) without and with the presence of oil shocks. The panel cointegration results for the two subgroups appear in Tables 3 and 4, respectively. The results in both groups of countries as well as in both cases match those reached for the full sample.

## 3. Conclusions

Our study re-examines the major implications of Lucas's misperception model using a longer time period and different methodology (a panel approach). The empirical analysis uses annual data from 41 countries over 1949 to 1999. It also permits the monetary regimes to change over the sample. Finally, the empirical analysis also specifies Lucas's model explicitly to consider the simultaneous effect of monetary and real (oil) shocks on output behavior.

We consider three hypotheses: Lucas's variance hypothesis-, (negative correlation between $lnx$ *and* $\sigma^2_{rm}$), the neutrality of anticipated money growth hypothesis, (no correlation between $lnx$ and $\mu_{\Delta m}$), and the neutrality of the variance of anticipated money growth hypothesis, (no correlation between $ln$x and $\sigma^2_{\Delta m}$). Our panel results support the first two hypotheses both without and with real oil price shocks.[10] The robustness of our results remains extremely high when we also split our full sample into a developed-countries sample and a less-developed-country sample.

Our approach allows the inclusion of oil price shocks to play a specific role in determining the behavior of output. We find that Lucas's propositions receive support from the data. Our results contradict those reached by Poirier (1991), but mirror closely those reached by K-M (1984, 1991).

**Appendix: Division of Countries in the Sample**

We collect annual data on money (IFS series 34), real GDP (IFS series 99b.p or 99b.r) at 1990 prices, and CPI (IFS series 64) were employed for the following countries:

*Developed countries:*

Australia, Austria, Belgium, Canada, Denmark, Finland, France, Germany, Greece, Icelend, Ireland, Israel, Italy, Japan, Netherlands, New Zealand, Norway, Portugal, Spain, Sweden, Switzerland, Turkey, UK, and US.

*Less developed countries:*

Colombia, Dominican Republic, Equador, El Salvador, Guatemala, Honduras, India, Korea, Mexico, Paraguay, Philippines, South Africa, Sri Lanka, Taiwan, Thailand, Uruguay, and Venezuela.

For Mexico, GDP is in current prices (IFS series 99b) and real GDP equals current GDP divided by the CPI series (1990=100). We did not find reliable data for Brazil, Burma, Chile, and Nigaragua because of non-available values in certain years. Thus, we did not include those countries in our sample. Parantap Basu (Fordham University) kindly provided data for India, Korea, and Taiwan. Finally, in all of our estimations, a dummy variable for the UK captures a surge in reported M1 for the UK in 1989, stemming from the re-incorporation of the UK's largest building society into a bank.

Finally, we define nominal oil prices as international petroleum average spot crude prices as published by the IMF.

**Notes**

1. Friedman (1968) and Phelps (1968) tell stories about information problems that generate results consistent with the Lucas supply curve.
2. Barro and King (1984), Barro (1989a), and Lucas (1996) distance themselves from the imperfect information model. In particular, they argue that money surprises leave output unchanged, while a positive relationship between money and output may reflect the endogenous response of money (Barro, 1989b). The supporters of the misperceptions model now focus their attention on analyses where real shocks mainly determine business fluctuations [i.e., real business cycle (RBC) theory]. Chatterjee (1999) argues, however, that empirical evidence shows that money still matters for business fluctuations, because better monetary control has, among others, reduced GNP volatility in the U.S. Our paper, however, does not pursue the importance of monetary versus real factors in explaining business fluctuations.
3. Froyen and Waud (1984), however, provided limited intra-country evidence against the negative relationship just described.
4. Lucas (1973) focuses on the aggregate-relative confusion between firms. Friedman (1968) and Phelps (1968) emphasize the confusion between workers and firms.
5. We do not report those findings, but will make them available on request.
6. We also perform similar analyses where the monetary regimes emerge from estimating recursive single-equation ARIMA models for the money supply growth rates. The findings match those reported in the text. We will make those findings available on request.
7. According to Kaul (1990) and Hakes and Gamber (1992), certain monetary regimes changes occurred over the post-war period. Such paradigms are: the 1961-1979 regime in

which the Fed focus on reducing interest rate variability, the 1975-1979 federal funds rate control regime, the 1979-1982 non-borrowed reserves control regime, and the 1982-1987 borrowed reserves control regime for the US; the 1951-1960 regime in which the monetary authorities regulated the total quantity of money and the 1961-1983 credit control regime in Canada; the 1957-1967 interest rate control regime and the 1968-1992 money supply control regime in Germany. We do not have explicit monetary regime changes for all countries in the sample, but we have assumed that monetary actions occurring in large developed countries, i.e. US, Canada, and Germany, could have easily affected those conditions in small developed and less developed countries.

8. We do not report those results, but make them available upon request. The empirical task identifies anticipated and unanticipated money growth rates. Thus, the key point revolves around generating the 'best' one-step-ahead forecast of the money supply growth rate. We employ the bivariate VECVAR in the text. We also repeat the analysis for univariate ARIMA specifications as an alternative. We do not report the findings when using the ARIMA alternative, but also make them available on request.
9. Also available upon request.
10. A referee noted that a substantial component of monetary variability probably relates mainly to the variability of the corresponding shocks, which evidently produces the conclusion that expected monetary variability closely links with output growth in the long run.

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**TABLE 1: Panel Unit Root Tests**

| Variable | Without Trend | With Trend |
|---|---|---|
| *(All countries)* | | |
| $m$ | -0.43(5) | -1.38(3) |
| $\Delta m$ | -7.63(4)* | -9.38(3)* |
| $y$ | -1.29(4) | -1.32(6) |
| $\Delta y$ | -7.17(3)* | -10.24(4)* |
| $op$ | -0.78(5) | -1.39(6) |
| $\Delta op$ | -8.52(4)* | -9.85(5)* |
| $lnx$ | -0.95(4) | -1.07(4) |
| $\Delta lnx$ | -6.19(3)* | -6.77(2)* |
| $\sigma^2_{rm}$ | -1.54(3) | -1.71(3) |
| $\Delta\sigma^2_{rm}$ | -4.96(2)* | -4.82(2)* |
| $\mu_{\Delta m}$ | -1.15(3) | -1.57(3) |
| $\Delta\mu_{\Delta m}$ | -5.85(2)* | -5.28(2)* |
| $\sigma^2_{\Delta m}$ | -1.29(4) | -1.42(4) |
| $\Delta\sigma^2_{\Delta m}$ | -4.74(2)* | -5.01(3)* |
| *(Developed countries)* | | |
| $m$ | -0.65(4) | -0.77(3) |
| $\Delta m$ | -5.78(4)* | -6.12(2)* |
| $y$ | -0.99(3) | -1.35(4) |
| $\Delta y$ | -6.32(2)* | -7.37(4)* |
| $op$ | -0.61(5) | -1.11(5) |
| $\Delta op$ | -6.39(3)* | -7.19(3)* |
| $lnx$ | -1.14(4) | -1.41(3) |
| $\Delta lnx$ | -5.72(3)* | -5.89(2)* |
| $\sigma^2_{rm}$ | -1.38(3) | -1.55(3) |
| $\Delta\sigma^2_{rm}$ | -6.18(2)* | -6.71(2)* |
| $\mu_{\Delta m}$ | -1.03(4) | -1.22(3) |
| $\Delta\mu_{\Delta m}$ | -4.95(3)* | -5.77(2)* |
| $\sigma^2_{\Delta m}$ | -1.13(4) | -1.25(4) |
| $\Delta\sigma^2_{\Delta m}$ | -5.38(2)* | -5.81(2)* |

**TABLE 1: Panel Unit Root Tests (continued)**

| **Variable** | **Without Trend** | **With Trend** |
|---|---|---|
| *(Less developed countries)* | | |
| $m$ | -0.61(4) | -1.68(3) |
| $\Delta m$ | -4.19(3)* | -6.19(2)* |
| $y$ | -1.04(4) | -1.79(5) |
| $\Delta y$ | -4.93(4)* | -6.56(3)* |
| $op$ | -0.81(3) | -1.83(4) |
| $\Delta op$ | -5.02(3)* | -6.77(4)* |
| $lnx$ | -1.26(4) | -1.53(4) |
| $\Delta lnx$ | -5.28(3)* | -5.92(2)* |
| $\sigma^2_{rm}$ | -1.18(3) | -1.47(3) |
| $\Delta\sigma^2_{rm}$ | -4.59(2)* | -4.76(2)* |
| $\mu_{\Delta m}$ | -1.14(3) | -1.35(3) |
| $\Delta\mu_{\Delta m}$ | -4.59(2)* | -5.22(2)* |
| $\sigma^2_{\Delta m}$ | -1.44(3) | -1.83(3) |
| $\Delta\sigma^2_{\Delta m}$ | -4.88(2)* | -5.12(2)* |

**Notes:** Figures in brackets denote the number of lags in the augmented term that ensures white-noise residuals.

* significant at the 1% level.

**TABLE 2: Panel Cointegration Tests (Pedroni's Tests) - Full Country Sample**

***The Lucas Variance Hypothesis: Long-Run Negative Relationship between lnx and $\sigma^2_{rm}$***

| | (Without oil shocks) | (With oil shocks) |
|---|---|---|
| Panel v-stat | -3.2284* | -3.6762* |
| Panel rho-stat | -3.0842* | -3.1099* |
| Panel pp-stat | -3.1288* | -3.7355* |
| Panel adf-stat | -3.8549* | -3.6921* |
| Group rho-stat | -3.8042* | -3.1209* |
| Group pp-stat | -3.9477* | -3.8935* |
| Group adf-stat | -3.3096* | -3.8561* |

***The Neutrality of Anticipated Money Growth: No Long-Run Correlation between lnx and $\mu_{\Delta m}$***

| | (Without oil shocks) | (With oil shocks) |
|---|---|---|
| Panel v-stat | -0.1344 | -0.1498 |
| Panel rho-stat | -0.2488 | -0.1238 |
| Panel pp-stat | -0.1984 | -1.3374 |
| Panel adf-stat | -0.5094 | -0.5699 |
| Group rho-stat | -0.3106 | -0.4095 |
| Group pp-stat | -0.4233 | -1.5582 |
| Group adf-stat | -0.9005 | -0.6892 |

***The Neutrality of the Variance of Anticipated Money Growth: No Long-Run Correlation between lnx and $\sigma^2_{\Delta m}$***

| | (Without oil shocks) | (With oil shocks) |
|---|---|---|
| Panel v-stat | -4.3566* | -4.6106* |
| Panel rho-stat | -4.6940* | -4.8459* |
| Panel pp-stat | -4.5904* | -4.4881* |
| Panel adf-stat | -4.7105* | -3.1095* |
| Group rho-stat | -3.6783* | -3.7792* |
| Group pp-stat | -3.0119* | -3.4894* |
| Group adf-stat | -3.9236* | -3.0096* |

**Notes:**

* Rejection of the null hypothesis of no cointegration at the 1% level.

**TABLE 3: Panel Cointegration Tests (Pedroni's Tests) - Developed Country Sample**

***The Lucas Variance Hypothesis: Long-Run Negative Relationship between lnx and $\sigma^2_{rm}$***

| | *(Without oil shocks)* | *(With oil shocks)* |
|---|---|---|
| Panel v-stat | -3.8492* | -3.9973* |
| Panel rho-stat | -3.4781* | -4.1983* |
| Panel pp-stat | -3.1129* | -3.7904* |
| Panel adf-stat | -4.0955* | -3.8992* |
| Group rho-stat | -3.8749* | -3.6582* |
| Group pp-stat | -3.5812* | -3.7564* |
| Group adf-stat | -3.7181* | -3.9337* |

***The Neutrality of Anticipated Money Growth: No Long-Run Correlation between lnx and $\mu_{\Delta m}$***

| | *(Without oil shocks)* | *(With oil shocks)* |
|---|---|---|
| Panel v-stat | -0.1459 | -0.2774 |
| Panel rho-stat | -0.2891 | -0.3391 |
| Panel pp-stat | -0.1783 | -0.4708 |
| Panel adf-stat | -0.5193 | -0.5506 |
| Group rho-stat | -0.3094 | -0.5038 |
| Group pp-stat | -0.3982 | -0.5122 |
| Group adf-stat | -0.6892 | -0.6991 |

***The Neutrality of the Variance of Anticipated Money Growth: No Long-Run Correlation between lnx and $\sigma^2_{\Delta m}$***

| | *(Without oil shocks)* | *(With oil shocks)* |
|---|---|---|
| Panel v-stat | -4.2904* | -4.4788* |
| Panel rho-stat | -4.5044* | -4.6373* |
| Panel pp-stat | -4.4571* | -4.5180* |
| Panel adf-stat | -4.3292* | -4.4066* |
| Group rho-stat | -3.6551* | -3.8871* |
| Group pp-stat | -3.5938* | -3.6930* |
| Group adf-stat | -3.7103* | -3.8404* |

**Notes:** See Table 2.

**TABLE 4: Panel Cointegration Tests (Pedroni's Tests) - Less Developed Country Sample**

| ***The Lucas Variance Hypothesis: Long-Run Negative Relationship between lnx and $\sigma^2_{rm}$*** | | |
|---|---|---|
| | *(Without oil shocks)* | *(With oil shocks)* |
| Panel v-stat | -3.5381* | -4.2883* |
| Panel rho-stat | -3.6260* | -3.9455* |
| Panel pp-stat | -3.7559* | -3.8109* |
| Panel adf-stat | -3.7233* | -3.8832* |
| Group rho-stat | -3.8832* | -3.8905* |
| Group pp-stat | -3.8047* | -4.0466* |
| Group adf-stat | -3.5005* | -3.8109* |
| ***The Neutrality of Anticipated Money Growth: No Long-Run Correlation between lnx and $\mu_{\Delta m}$*** | | |
| | *(Without oil shocks)* | *(With oil shocks)* |
| Panel v-stat | -0.1772 | -0.3391 |
| Panel rho-stat | -0.3009 | -0.3692 |
| Panel pp-stat | -0.4128 | -0.4692 |
| Panel adf-stat | -0.5229 | -0.7710 |
| Group rho-stat | -0.5703 | -0.6626 |
| Group pp-stat | -0.6107 | -0.7199 |
| Group adf-stat | -0.6725 | -0.7452 |
| ***The Neutrality of the Variance of Anticipated Money Growth: No Long-Run Correlation between lnx and $\sigma^2_{\Delta m}$*** | | |
| | *(Without oil shocks)* | *(With oil shocks)* |
| Panel v-stat | -4.3372* | -4.4720* |
| Panel rho-stat | -3.7932* | -3.8301* |
| Panel pp-stat | -3.6081* | -3.6683* |
| Panel adf-stat | -3.5532* | -3.6820* |
| Group rho-stat | -3.7449* | -3.7719* |
| Group pp-stat | -3.8201* | -3.9347* |
| Group adf-stat | -4.3962* | -4.4981* |

**Notes:** See Table 2.