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The Dominant Influence of Fiscal Actions in Developing Countries

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1. INTRODUCTION

The relative efficacy of monetary and fiscal policy in economic stablization is a matter of debate among economists and policy-makers. Empirical research has focused mainly on the experience of the United States using the so called "St. Louis single-equation" approach (see, for example 2,3,8,9,28). The findings which have emerged seem to suggest that monetary actions have a stronger, more predictable and faster impact on economic activity in the United States than fiscal actions.[1] Evidence on the relative superiority of monetary actions has been also advanced by some empirical studies in the case of other developed economies (see, for example, 5,14,31). However, such persistent results for several developed economies having roughly similar economic structures may not be generalized for developing countries with significantly different economic and financial environments. Study of the relative impact of monetary and fiscal impulses on economic activity has been largely neglected in the case of developing economies.[2]

This paper is intended to fill this gap. It examines empirically, within a modified St. Louis single-equation approach, the relative importance of monetary and fiscal actions in determining economic activity in five major Latin American countries, namely Brazil, Chile, Mexico, Peru and Venezuela over the period 1950 through 1981.[3] It is hoped that subjecting the St. Louis model to empirical testing with data from developing economies would further assess the empirical usefulness of this single-equation approach for analyzing the relative impacts of monetary and fiscal actions, and provide additional evidence on the robustness of the approach across countries with a variety of economic structures.

While the original St. Louis equation forms the basis of the empirical estimations in this study, we have introduced some modifications and employed several statistical tests. First, significant heteroscedasticity problem has led to the use of an alternative (growth-rate) version of the model instead of the original and more common first-difference format. Second, given that the developing countries in our sample are open economies, a measure of the external trade influence was included in the estimated model. Third, due to severe criticisms and limitations, the popular Almon distributed-lag estimation method was not employed. Rather, we used unconstrained ordinary least-squares procedure in which the appropriate lag specifications underlying the model for each country were determined on the basis of Hsiao's 29 multivariate technique.

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Fourth, the issue of statistical exogeneity is explored, using Granger-type causality tests. Fifth, the question of temporal stability of the estimated regressions is also examined by utilizing a battery of alternative stability tests.

This paper is organized as follows. Section 2 briefly outlines the model to be estimated and discusses various data problems encountered in the empirical estimations. Section 3 reports and interprets the empirical results obtained for the five Latin American countries. Section 4 extends the empirical investigation and tests for the implied exogeneity assumptions and explores the stability properties of the estimated relationships. A closing section provides a summary and draws some conclusions.

2. THE ESTIMATED MODEL

The St. Louis single-equation model originally countained three main variables: a measure of economic activity as the dependent variable, and two independent variables which serve as a measure of monetary and fiscal actions.[4] Typically, nominal Gross National Product has been used as the measure of economic activity, while some definitions of the money stock and government expenditures have been employed as indicators of monetary and fiscal impulses respectively.

The above specification of the St. Louis equation appears inappropriate for developing countries whose economies are largely influenced by foreign economic developments, since it implicitly assumes that the economy under study is closed. Consequently we have included exports as an additional explanatory variable in explaining nominal income in each of the Latin American countries.[5] The modified St. Louis equation can thus be written as

$$Y = f(M, G, S) \qquad f_1, f_2, f_3 > 0 \tag{1}$$

where Y, M, G and S are respectively nominal GNP, the money stock, government spending, and exports. A priori, we expect that nominal GNP responds positively to changes in the three explanatory variables. To make equation (1) operational, two additional analytical issues need to be resolved.

First, the proper definitions of M and G need to be specified. For the money stock variable, we employed the narrow definition of money stock (currency plus demand deposits). This definition was chosen in order to use a comparable and consistent definition of the monetary variable across all countries studied. As regards the government spending variable, we utilized total government spending (including transfer payments and purchases of goods and services) as the measure of the fiscal variable.[6]

Second, the mathematical form of equation (1) must be established. Since economic theory provides no rationale as to what is correct, we follow the original practice and estimate the equation for each country using the arithmetic first-difference format (with a constant). This is convenient in that it provides direct estimates of the relevant multipliers. However, applying the Goldfeld-Quandt [22] test, it was found that the regressions are beset by the problem of heteroscedasticity (unconstancy of error variances across all observations). With heteroscedastic error terms, the standard t and F tests become invalid and thus no correct inference can be made about the significance of the estimates. A proper solution to this statistical problem is to transfer the model so as to satisfy the requirement of homoscedastic error terms. Upon testing, the growth-rate (logarithmic first-difference) was chosen as the mathematical form of our basic equation.[7] Therefore, equation (1) can now be written in the following estimatable form (the dots over the variables indicate growth rates):

$$\dot{Y}_t = a_0 + \sum_{i=0}^{\ell_1} m_i \dot{M}_{t-i} + \sum_{i=0}^{\ell_2} g_i \dot{G}_{t-1} + \sum_{i=0}^{\ell_3} s_i \dot{S}_{t-i} + \varepsilon_t \tag{2}$$

where all the variables are defined as before, a_0, m_i, g_i and s_i are the coefficients to b estimated[8], and is the error term which, in the usual fashion, is assumed serially uncorrelated and normally distributed with zero mean and constant variance. Past as well as current values of the explanatory variables have been included in equation (2) in recognition of the fact that current nominal GNP growth may also respond to lagged changes in the relevant explanatory variables. A priori, the theory provides us with the following expected signs for the sum coefficients:

$$\Sigma m_i > 0, \ \Sigma g_i > 0, \text{ and } \Sigma s_i > 0$$

Equation (2) represents our basic model; estimates are obtained for the five Latin American countries using the annual data over the period 1950 through 1981. Before reporting our empirical results in the next section, a comment about the method of estimation is in order. The St. Louis equation is usually estimated using the Almon distributed-lag technique which forces the coefficients of each lagged explanatory variable to lie on a polynomial function of a certain degree chosen a priori by the researcher. However, this popular technique has increasingly come under attack because of its limitations and possible specification biases.[9] Consequently, equation (2) has been estimated by unconstrained ordinary least squares method. For each explanatory variable, the optimal lag lengths were determined on the basis of Hsiaoi's 29 procedure[10] as well as Theil's 40 residual-variance criterion.[11]

3. THE EMPIRICAL RESULTS

Table 1 reports the regression results from estimating equation (2) for the selected group of the Latin American countries. Based on Hsiao's and Theil's criteria, the particular lag lengths in Table 1 gave the best empirical results. The high values of and the low values of S.E. shown in Table 1, indicate that the modified St. Louis model exhibits consistently excellent explanatory power with respect to GNP growth across different economies.[12] After correcting for first-order serial correlation, the regressions generally appear to be free from remaining serial correlation in the residuals according to the scores of the Durbin-Watson statistics.[13] Furthermore, the Goldfeld-Quandt test suggests that the regressions (in growth rate format) do not suffer from any significant heteroscedasticity problem.

As to the estimates of the coefficients, the results confirm the hypothesis that export growth is an important argument in explaining GNP growth. The variable appears with the expected positive sign and with statistically significant coefficient at the conventional 5 percent level of significance in all countries examined except for Chile where the coefficient is statistically significant at the 10 percent level.[14] The impact of export impulse, although generally smaller than that of the monetary and fiscal variables, is significally non-zero in all cases ranging in magnitudes from 0.11 for Brazil to 0.31 for Peru.

The regression results presented in Table 1 also yield substantially consistent implications with respect to fiscal actions. In every Latin American country examined in

TABLE 1: Regression Results from the Modified St. Louis Equation For Some Latin American Countries, Annual Data: 1950-1981 (in Millions of National Currency Units)

$$\dot{Y}_t = a_0 + \sum_{i=0}^{\ell_1} m_i \dot{M}_{t-i} + \sum_{i=0}^{\ell_2} g_i \dot{G}_{t-i} + \sum_{i=0}^{\ell_3} s_i \dot{S}_{t-i} + \varepsilon_t$$

	Brazil	Chile	Mexico	Peru	Venezuela
Constant	-.144 (3.30)	-.028 (.63)	.043 (2.13)	-.036 (1.76)	.043 (4.21)
m_0	.362 (4.42)	.117 (.69)	.093 (2.48)	.154 (1.58)	.109 (1.24)
m_1	.177 (2.04)	-.324 (.99)	-.037 (.60)	.167 (1.80)	
m_2		.423 (1.92)	-.023 (.31)	.049 (.54)	
m_3			.059 (.94)	.127 (1.53)	
m_4			.047 (1.11)		
Σm_i	.539 (4.06)	.216 (.58)	.140 (.62)	.50 (2.95)	.109 (1.24)
g_0	.381 (5.11)	.675 (2.70)	.280 (3.53)	.060 (.68)	.182 (2.81)
g_1	.127 (2.17)			.172 (1.76)	.023 (.36)
g_2	.134 (2.08)			.018 (.21)	
g_3	.156 (2.96)			.180 (2.13)	
Σg_i	.798 (4.71)	.675 (2.70)	.280 (3.53)	.430 (2.96)	.205 (2.89)
s_0	.109 (2.16)	.224 (4.63)	.069 (1.50)	.183 (4.59)	.285 (7.45)
s_1		.061 (1.25)	.128 (3.17)	.054 (1.19)	
s_2		-.089 (1.54)	.023 (.56)	.076 (1.49)	
s_3		-.070 (1.51)	.089 (2.33)		
Σs_i	.109 (2.16)	.126 (1.38)	.309 (3.02)	.313 (4.14)	.285 (7.45)
$\bar{R}^2$	.916	.969	.854	.944	.844
F	30.881	92.021	10.442	31.071	33.230
S.E.	.03177	.10523	.02268	.03070	.03467
RHO	.471	-.226	.368	.137	-.141
D.W.	1.68	1.99	2.07	1.91	1.99
G.Q.	.44	1.02	.48	1.05	1.04

Notes to Table 1

The numbers in parentheses are absolute values of t-statistics. $\bar{R}^2$ is the coefficient of multiple determination adjusted for degrees of freedom; F-value is for testing the null hypothesis that all the right-hand side variables as a group except the constant term have zero coefficient; S.E. is the standard-error of the regression; RHO is the estimate of the first-order serial correlation adjustment coefficient used by the Beach-Mackinnon maximum likelihood procedure to correct for serial correlation; D.W. is the Durbin-Watson statistic to test for remaining serial correlation; and G.Q. is the Goldfeld-Qundt F-ratios to test for homoscedasticity.

this paper, the cumulative impact of government spending growth is positive and significantly non-zero at the 5 percent level of significance. In contrast, the cumulative impact of money growth, while positive in every country, is significantly non-zero at the 5 percent level in only two cases (Brazil and Peru). These results for the Latin American countries are interesting because they seem to imply that, contrary to the case of developed countries, in the Latin American countries the behavior of nominal income reflects the dominance of fiscal rather than monetary actions. Because of the critical implications of the issues involved, these results are further analyzed below.

In earlier debates over the relative impact of monetary and fiscal actions on economic activity in developed countries, three propositions were commonly tested. These propositions considered whether monetary or fiscal actions have impacts that are (1) stronger, (2) more predictable, and (3) faster-acting. The frequently reached conclusion was that monetary actions dominate fiscal actions in each proposition. We will now examine these same propositions in the light of the experience of the five Latin American economies.

To make comparable examination of the relative strengths of monetary and fiscal total impact on economic activity, the estimated sum coefficients are normalized for each country by converting them into beta summed coefficients which are displayed in Table 2.[15] The beta summed coefficient for the fiscal measure is considerably larger than that for the monetary measure for every country. Moreover, although the beta summed coefficients are of the appropriate positive signs for both measures, it is with respect to fiscal actions that these coefficients are consistently significantly non-zero across all countries. The beta summed impact of the monetary measure, on the other hand, is significantly non-zero only in Brazil and Peru. Clearly, then, over the 30-year sample period, fiscal actions have exerted significantly a stronger influence on economic activity than have monetary actions in the case of the Latin American developing countries.

TABLE 2
Calculated Beta of the Sum Coefficients

Country	Monetary Influences	Fiscal Influences
Brazil	.436*	.760*
Chile	.161	.636*
Mexico	.203	.382*
Peru	.443*	.488*
Venezuela	.125	.324*

Notes to Table 2

For either policy variable, the beta of the sum coefficient is defined as the estimated summed coefficient times the ratio of the standard deviation of that variable over the standard deviation of the dependent variable.

* indicates significance at the 5 percent level.

As to proposition (2) above, one commonly used indicator of the relative predictability of monetary and fiscal impacts on nominal income is the relative size of the t-statistics of the corresponding sum coefficients (see, for instance, 30,31). It is argued that the larger the t-value, the more confident we may become that the "true relationship" between nominal income and monetary or fiscal actions has the same sign as that of the statistically estimated relationship between those variables. As Table 1 shows, in every country, the t-value for the fiscal summed coefficient is uniformly larger than the t-value for the monetary summed coefficient. A related feature of the results can also be distilled by estimating two alternative equations that isolate the relative explanatory power of the monetary and fiscal variables in explaining movements in GNP growth. Thus, for each country, equation (2) was re-estimated once without the fiscal variable, and then in a second set of estimations, we dropped the monetary variable instead. To economize on space, the detailed regression results from these alternative specifications are not reported here. However, Table 3 reports the calculated F-statistics to test for the significance of the improvement in the fit gained due to the inclusion of either variable. In every country studied, once the influence of the fiscal actions is taken into account, the overall explanatory power of the equation is not significantly improved by the inclusion of the monetary variable. On the other hand, except for Venezuela, the addition of the fiscal variable to the equation which has already included the monetary variable does significantly improve the overall explanatory power of the equation. These results further point to the statistical dominance of government spending growth over monetary growth in explaining movements in GNP growth, and that the GNP-government spending link is generally more robust than the GNP-money link. It can, therefore, be argued that for our group of Latin American countries fiscal actions appear to have had consistently more predictable and more robust impact on economic activity than have monetary actions over the sample period.

TABLE 3
Calculated F-Statistics for Testing the Significance of the Relative Explanatory Powers of Monetary and Fiscal Influences

Country	Contribution in Explanatory Power Due to Monetary Influences	Contribution in Explanatory Due to Fiscal Influences
Brazil	.30	4.27*
Chile	.32	16.68*
Mexico	.28	6.80*
Peru	1.87	3.48*
Venzuela	1.99	.95

Note to Table 3
*indicates significance at the 5 percent level.

The final proposition tested concerns the relative speed with which monetary and fiscal actions exert their impacts on nominal income. This aspect can be addressed by observing which policy measure has the shorter time lag in affecting nominal income. In order to assure comparable results, the annual patterns of the estimated beta coefficients are examined. Table 4 reports the percentage of the beta summed coefficients occuring within the first year of the policy change. In general, the impact of government-spending growth on GNP growth is more rapid than that of monetary growth. Indeed, in the case of Chile and Mexico, the impact of government-spending growth is fully completed within the firt year of the change. The lag patterns for the monetary variable, in contrast, do not compare as well, except for Brazil and Venezuela. Thus, for at least three out of five of the countries examined, fiscal actions tend to exert a faster impact on nominal income than do monetary actions.

TABLE 4
Percentage of Beta Sum Coefficients Occuring Within the First Year of the Policy Change

Country	Monetary Influence	Fiscal Influences
Brazil	67	48
Chile	54	100
Mexico	66	100
Peru	31	40
Venezuela	100	89

4. CAUSALITY IMPLICATIONS AND STABILITY PROPERTIES OF THE ESTIMATED RELATIONSHIPS

The econometric validity of our reported empirical results depends crucially on the assumption that the right-hand side variables in equation (2) are exogenous in the statistical sense. Violation of this basic assumption leads to single-equation estimates that are both biased and inconsistent. Furthermore, the practical usefulness of these empirical results for policy analysis and formulation hinges critically on the statistical stability of the estimated regressions. Structurally unstable relationships render the obtained empirical results virtually useless for forecasting and policy purposes. Consequently, these two statistical aspects of our results are now explored.

To test for statistical exogeneity, we employed the procedure proposed by Sargent 35 to test for causality in the sense of Granger 25 .[16] A priori, it can be argued that, except for the monetary growth variable, the other right-hand side variables can be considered statistically exogenous to GNP growth in equation (2). Following Keran 30 among others, government expenditures are assumed to be primarily determined by the fiscal authorities rather by the spending behavior of the public. On the other hand, as demonstrated theoretically by Turnovsky 43 , exports are basically determined by foreign rather than by domestic income and by the ratio between foreign and domestic prices. Because econometric technique can not, in fact, substitute for economic theory to determine causality, we will assume that the growth of government-spending and exports are statistically exogenous to GNP growth.[17] However, an equally strong rationale cannot be provided concerning the statistical exogeneity of the monetary growth variable with respect to GNP growth. Indeed, economic theory suggests that both the public as well as the monetary authority participate in the determination of the monetary stock. Consequently, we will utilize the Granger test of causality to test for the statistical exogeneity of monetary growth to GNP growth.[18] The Granger test results for each Latin American country are reported in Table 5.[19]

The calculated F-statistics indicate that, for Brazil, Chile and Venezuela, the monetary growth variable is statistically exogenous to GNP growth at the 5 percent level of significance. The empirical results for the remaining two countries are also interesting. For these two countries (Mexico and Peru), the calculated F-statistics suggest that monetary growth and GNP growth are statistically independent. Keeping in mind the various caveats associated with testing for causality, two interesting findings implied by these results. First, given the above evidence, it may be argued that the regression results presented in this study appear econometrically valid insofar as they do

TABLE 5
Calculated F-Statistics for the Granger-Causality Test

Country	Null Hypothesis: Money Growth Does Not Granger--Cause GNP Growth	Null Hypothesis: GNP Growth Does Not Granger--Cause Money Growth
Brazil	18.57*	4.41
Chile	11.07*	2.15
Mexico	1.66	1.01
Peru	2.04	2.19
Venezuela	5.44*	3.35

Note to Table 5
* indicates rejection of the null hypothesis at the 5 percent level of significance.

not suffer from significant simultaneous-equation bias. Second, the results also show that for two countries of our sample (Mexico and Peru), there seems to be no reliable statistical relationship between monetary growth and GNP growth. This latter finding further supports the previous results which point to the statistical dominance of fiscal actions over monetary actions in explaining movements in GNP growth.

Finally, as to testing the temporal stability of the estimated equations, three alternative techniques were employed, namely the Chow 10 , the Farley-Hinich 15 , and the Gujarati 27 tests. Because different stability techniques explore different types of stability, alternative stability tests were utilized to generate sufficient evidence on the stability property of the estimated relationships.[20] To apply the Chow and the Gujarati tests, the sample period was split at the mid-point in order to maximize the power of the tests as suggested by Farley, Hinich and McGuire 16 . The empirical results from all three tests are reported in Table 6. Except for the Chow test and then only in the case of Brazil, the three tests clearly suggest that the estimated equations of Table 1 for the five Latin American countries are all structurally stable at the 5 percent level of significance.[21] Given this systematic evidence of structural stability over time, it can thus be argued that the estimated regressions presented in this study are useful and reliable for policy analysis and formulation.

5. SUMMARY AND CONCLUSION

This paper has used a modified St. Louis equation to investigate the relative impact of monetary and fiscal actions on economic activity in the light of the experience of five Latin American countries during the past three decades. Statistically, the regression results obtained are very satisfactory with respect to demonstrating that the St. Louis equation-type is an appropriate framework of analysis not only for developed countries, for which the model has been restrictively applied, but also for several developing countries which have significantly different economic environments. Moreover, the regression estimates of the modified St. Louis model appear econometrically valid insofar as they do not suffer from significant simultaneity bias. In addition, the estimated equations for all countries examined are found to exhibit structural stabiliuty over time according to a battery of stability tests.

The results obtained from the modified St. Louis equation for the five Latin American countries suggest that export growth is an important argument in explaining movements in GNP growth which is an evidence for the economic openness of these countries. Perhaps more importantly, the results presented in this paper consistently suggest that fiscal actions substantially dominate monetary actions in explaining changes

in nominal income. Specifically, monetary policy measures (1) do not exert significant influences on GNP growth in three out of the five countries examined (Chile, Mexico, and Venezuela); (2) have no predictable impact on GNP growth in these three countries, and further do not significantly contribute to the explanation of changes in GNP growth in all five cases; (3) have somewhat faster impact on GNP growth than do fiscal policy measures only in two cases (Brazil, and Venezuela); and (4) bear no reliable relationship with GNP growth (the two variables are statistically independent) in two countries (Mexico and Peru).

In contrast, the fiscal policy measures (1)' exert significant influence on GNP growth in all countries examined; (2)' have more predictable impact on GNP growth in all countries examined, and significantly contribute to the explanation of changes in GNP growth in all countries (except for Venezuela); (3)' exert considerably faster impact on GNP growth than do monetary policy measures in three of the countries studied (Chile, Mexico and Peru); and (4)' bear statistically reliable relationship with GNP growth in all countries examined. The evidence presented in this study implies that fiscal actions are more effective for economic stabilization purposes than monetary actions in the case of the Latin American developing economies.

Footnotes

1. Predictably, these findings have provoked a number of counter studies. Among others, see 7, 13, 18, 24 .

2. To my knowledge, the only exception is the brief study by Atesoglu (1975) for the case of Turkey. Our approach differs from Atesoglu's at least in that a) we use the less controversial St. Louis model rather than the naive Friedman-Meiselman's earlier framework; b) we employ data over a 30-year period drawn from five Latin American countries; c) we attempt to correctly specify the underlying lag structures; d) we explicitly examine the exogeneity assumptions of the estimated model; and e) we investigate the temporal stability of the estimated relationships.

3. Lack of consistent and sufficiently long data available to this writer precluded the considertion of other Latin American countries in the present study. Furthermore, unavailability of quarterly data for two main variables (GNP and government spending) has necessitated the use of annual data.

4. As Batten and Hafer (1983) have pointed out, the equation is not designed to capture all of the exogenous variables that may affect economic activity. Under general assumptions, furthermore, such missing variables would not lead to biased results. In addition to be potentially free from significant specification errors, several other advantages of the single-equation approach have favored its use in this paper over the more complex structural-model approach. For a lucid discussion of the debate over the advantages and disadvantages of both approaches, (see 12, 30.)

5. Exports were also included in the equations estimated for some developed economies by Batten and Hafer (1983) and by Dewald and Marchon (1978). It should be noted that imports were not considered as an explanatory variable in order to avoid simultaneity bias since a priori imports are also influenced by domestic nominal income. Exports, however, are relatively free of such a problem. On this, see Turnovsky, 1977.

6. The main source of all data used in this study is various issues of International Financial Statistics published by International Monetary Fund. Where needed, data were also drawn from Economic Survey of Latin America and other official documents of the relevant Latin American governments.

7. The growth-rate format has also proved appropriate for examining the experience of some developed economies. (See, for instance, 5, 8.)

8. Note that these coefficients are the estimates of the corresponding elasticities.

9. For an account of the caveats associated with the Almon method see 20, 37, 41 .

10. For each explanatory variable, Hsiao's procedures amounts to choosing that particular lag length which minimizes Akaike's 1 final prediction error.

11. The residual-variance criterion has been recommended by a number of econometricians for choosing among alternative model specifications. For example, see 21, 37 .

12. Considering that the estimations were based on growth-rate format, such high values of (ranging from .84 to .97) are indeed very good.

13. For the case of Brazil, however, the D-W test is inconclusive.

14. An F-test, not shown here, has also indicated that the improvement in the fit gained from the inclusion of the export growth variable is significant at the 5 percent level in all cases.

15. For any explanatory variable, the beta of the summed coefficient is defined as the estimated summed coefficient times the ratio of the standard deviation of the explanatory variable over the standard deviation of the dependent variable.

16. For a survey of causality testing in general and for a discussion of the difficulty in empirically testing for causality, see 11, 17, 34, 44 . The Sargent test is preferred here over the popular Sims 38 test becausde, as Gordon 23 and Granger 26 have pointed out, Sims' test is relatively more biased and inferior to Sargent's.

17. In so doing, we follow the suggestion of several writers in the causality literature. In particular see 11 .

18. This aspect of the debate is commonly known as the "reverse causation" problem. Note that the Granger-causality tests are interpreted here as tests of statistical exogeneity rather than tests of the broader philosophical concept of causality. Such interpretation is advocated recently by Lucas and Sargent 32 , and by Sims 39 .

19. In generating these results, we used the particular lag structures of Table 1. To employ alternative lag specifications, it was felt, may involve some data mining. Furthermore, because the Granger test requires the error terms to be approximately whitenoise, we followed Sargent's 35 suggestion and included a linear time trend and a constant in all estimated regressions.

20. For more on this point, see 6 . While the Chow and the Gujarati tests examine whether the estimated function has undergone a single-point shift, the Farley-Hinich technique, on the other hand, tests for a continuous structural shift during the entire sample period.

21. Note that the Brazilian regression should not be deemed inherently unstable because of the Chow test results since the regression has passed alternative stability tests. In addition, the Chow test is plagued with several problems that would make this popular test probably the least credible of most other alternative tests. For a discussion of the various weaknesses of the Chow test, see for example 33, 36, 42 .

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