The Relationship Between Poverty, Economic Growth, and Inequality Revisited

by

Lonnie K. Stevans
Associate Professor
Zarb School of Business
Department of BCIS/QM
127 Hofstra University
Hempstead, NY 11549-1270
Email: acslks@hofstra.edu
Phone: (516) 463-5375
http://www.i2.i-2000.com/~acslks/lks.html

David N. Sessions
Associate Professor
Zarb School of Business
Department of BCIS/QM
127 Hofstra University
Hempstead, NY 11549-1270
Email: david@sessionsdp.com
Phone: (516) 463-5719

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Abstract

It has been shown in prior research that higher growth reduces poverty. Authors have also found that the effect of economic growth on changes in poverty has either diminished or remained unchanged over time, e.g., the 1980s economic expansion in the U.S. had no affect on poverty. Using a formal error-correction model, we find that increases in economic growth are significantly related to reductions in the poverty rate for all families. However, growth was found to have a more pronounced effect on poverty during the expansionary periods of the 1960s, 1970s, 1980s, and 1990s. Other findings include identification of determinants of the dynamic behavior of poverty rates both in the short and long-term.

Introduction

According to the U.S. Census Bureau statistics, poverty rates have declined precipitously since 1959.[1] The official poverty rate has decreased from 22.4 percent of families in poverty in 1959 to 12.1 percent in 2002. Prior research in this area of study has concluded that this large decline in poverty rates may in fact be due to increased economic growth (Plotnick and Skidmore (1975), Aaron (1967), Perl and Solnick (1971), and Adams (2002). Of course, this regression of poverty rates may have also resulted from other factors, such as increases in transfer payments, more efficient labor markets, or from auspicious developments in the areas of education or discrimination (Thornton, et. al., (1978)).

From a policy perspective, as long as there is a strong negative correlation between economic growth and poverty that is expected to last into the future, then there may be less need for government programs that are intended to specifically reduce poverty. Consequently, it is important to know whether past changes in the historical or long-run relationship between GDP and poverty will endure for the present and into the future.

Many of the early investigations into the "trickle-down" model of economic growth in the U.S. have confirmed that economic growth alleviates poverty by increasing employment and/or the real wage (Anderson (1964), Thornton, et al. (1978), and Hirsh (1980)). Since past empirical studies have shown that growth and income inequality are

[1] http://www.census.gov/hhes/income/histinc/histpovtb.html

not related, sustained economic growth should have a large or more than proportionate effect on the poverty rate by raising everyone's income including the poor.[2]

One of the discoveries of current research on this theme is that the economic expansion of the 1980s did not reduce poverty significantly. Both Blank (1993) and Formby, et al. (2001) found that poverty in the 1960s was more responsive to economic expansion than it was in the 1980s. One explanation given was the sluggish growth of U.S. real wages in the 1980s—real wages of low-income workers rose by only one half of one percent during the 1980s expansion (Formby, et. al. (2001)). However, real wages can remain stagnant while real incomes increase due to improved employment opportunities. Real wages increased in the 1990s expansion by only one tenth of one percent, yet poverty declined by 3.7 percent.[3]

In this paper, we will investigate the dynamic relationship between economic growth and poverty in the context of a formal, error-correction model. While previous authors have estimated regressions using differenced variables, these model specifications have not allowed for the prospect of a "long-run" relationship among poverty and its time series' determinants. As mentioned previously, the presence of a long-run relationship affecting the short-run dynamics of poverty has important policy implications. Moreover, failure to consider this may have biased the results of previous studies, since the error-correcting term was omitted. We will also include a measure of income inequality as a determinant of poverty. In addition to analyzing the expansionary

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2 The empirical studies will be presented in the next section.
3 http://www.haverselect.com

periods of the 1960s and 1980s, as prior studies have done, we will also include the 1970s and 1990s. The time period utilized for this analysis is from 1959 to 1999.

Literature Review and Past Model Specifications

Do the poor benefit from economic growth? On one side, it is argued that the potential effect of economic growth on poverty rates is offset either entirely or in part by an increase in income inequality. Alternatively, economic growth may reduce poverty by raising the incomes of everyone in society, including the poor (Dollar and Kraay (2001)). The former assertion finds its roots in the Kuznets' hypothesis (1955), positing that growth and inequality are related through an inverted "U" shaped function. The implication here is that if economic growth leads to increased inequality, then the growth effect on poverty would be tenuous at best. The problem with the Kuznets' hypothesis (1955) is that the relationship between economic growth and inequality was derived from cross sectional data, e.g., using countries at different points of development at the same point in time, whereby what was really needed to test the hypothesis would be time series data (Adams (2002)). There are a number of empirical studies that have rejected the Kuznets' hypothesis (1955) (Ravallion (1995), Deininger and Squire (1998), Schultz (1998), and Bruno, et. al. (1998). There seems to be a general consensus that economic growth does not have much of an impact on inequality. Our investigation into this relationship utilizing more current econometric techniques yielded similar results, e.g., it was found that over the period 1959 to 1999, inequality, as measured by the Gini coefficient, was not cointegrated with Gross Domestic Product.[4]

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[4] For reasons of brevity, these results were not included in this study but will be made available upon request from the authors.

There is evidence that economic growth has reduced poverty in developing countries. Squire (1993) found that a ten percent increase in the growth rate reduced poverty across a sample of countries by 24 percentage points. In a similar study by Bruno, et. al. (1998), a ten percent increase in growth was associated with a 21.2 percent decrease in the poverty rate for a sample of 20 countries over the period 1984-1993. There is also evidence that changes in inequality affect changes in poverty, *ceteris paribus*. In the same aforementioned study by Bruno, et. al. (1998), the authors realized a statistically significant elasticity estimate of 3.86 on the inequality variable (Gini coefficient), leading them to conclude that "even small changes in the overall distribution of inequality can lead to sizeable changes in the incidence of poverty" (Bruno, et. al. (1978).

Using time series data, previous authors who conducted research on this topic utilized regression analysis with variables that were either first differenced or percent differenced, e.g., in difference form,

$$\Delta P_t = \beta_0 + \beta_1 \Delta GDP_t + \beta_2 \Delta U_t + \beta_3 \Delta TR_t + \beta_4 \Delta FEM_t + \beta_5 D \cdot \Delta GDP_t + \varepsilon_t, \qquad (1)$$

where,

ΔP_t - change in family poverty rate
ΔGDP_t - change in gross domestic product,
ΔU_t - change in the male unemployment rate,
ΔTR_t - change in transfer payments,
ΔFEM_t - change in number of female headed households,
D - dummy variable,
$D \cdot \Delta GDP$ - interaction term.

Multiple interaction terms may also be present. They are included in these types of models in order to ascertain whether the impact of growth on poverty is dissimilar for different years--the dummy variable would represent different time periods. Thornton, et al. (1978) found this interaction term to be negative and statistically significant in the post 1963 period relative to the 1947-1963 period—concluding, (in their words), that "trickle-down has petered-out." At least three of the above explanatory variables have been used in model specifications in past research.[5]

The specification of equation (1) is rather typical. Economic growth is calculated by the change or percentage change in GDP, while the other variables serve as controls. Thornton, et al. (1978) and Blank (1993) point-out that female-headed households demonstrate above average poverty rates, and increases in the male unemployment rate are undoubtedly associated with increases in poverty. The male unemployment rate is used because it is more stable than the overall rate (Formby, et. al. (2001). However, as argued by Blank (2000), there are conflicting results concerning the effect of transfer payments on poverty. According to Wallace and Blank (1999), transfer payments cannot only directly reduce poverty, but can cause those in poverty to find employment as there are reductions in welfare case loads. Alternatively, some authors such as Rector and Lauder (1995) believe in the welfare dependency hypothesis; that is, transfer programs can increase poverty by diminishing the incentives to look for and maintain gainful employment. The impact that transfers have on poverty is therefore theoretically ambiguous and becomes an empirical question.

[5] Thornton, et al. (1978) omits the female-headed households variable, while Formby, et al. (2001) includes it along with two interaction terms for the expansionary periods 1962–1972 and 1983-1989.

The use of only first differences or percent changes in the estimation of equation (1) has yielded some rather interesting results. Formby, et. al. (2001) estimated equation (1) over the period 1961 to 1996 (annual data) and utilized two interaction terms in order to determine whether the effect that economic growth had on changes in overall poverty was different during the economic expansions of the 1960s (1962-1972) and the 1980s (1983-1989), relative to the other years in the sample. The coefficient on the 1960s economic expansion interaction term was negative and statistically significant, while the coefficient on the 1980s interaction term was statistically insignificant. Thus, during the economic expansion of the 1960s, the effect of strong economic growth on changes in poverty was greater than it was in the 1980s and the result of robust economic growth in the 1980s on changes in poverty was not statistically different from zero. Formby, et. al. (2001) also found that neither a percentage change in transfer payments nor in the number of female-headed households had a statistically significant effect on changes in the poverty rate. These are interesting results, since it is well documented that a single-earner family is more likely to be in poverty than a multiple earner family. The authors' reason for the statistical insignificance of the number of female-headed households is what they believe to be the use of a wrong variable. Rather than the number of female-headed households, the number of *never-married* female heads was used, since they are a demographic group with the highest poverty rates. However, when utilizing this new variable, they still could not find a statistically significant relationship. As far as transfer payments are concerned, the authors (Formby, et. al. (2001)) simply stated that since the sign of the coefficient is positive but statistically insignificant, there is no support for the welfare dependency hypothesis. As will be mentioned shortly, it is our contention that these anomalous results reported by Formby, et. al. (2001) and others may be due to the fact that their models are miss-specified which could lead to biased coefficient estimates and inflated standard errors.

In a current paper by Enders and Hoover (2003), a threshold regression and a Fourier approximation model is fit to poverty-economic growth data.[6] They conclude that there is a large and significant effect on poverty as a result of the 1980s expansion—**which runs counter to the findings of previous studies on this subject**. The authors make the case that their specifications provide a better empirical model of poverty than what was done in prior research. However, they also fail to take into account the possibility of cointegrating relationships and as such, the resulting Fourier and threshold model specifications may also be miss-specified.

An Error-Correction Model

It is our contention that any empirical results that emanate from the estimation of equation (1) may be subject to misspecification error. What is missing is the notion of an equilibrium long-run relationship and the introduction of past disequilibrium as an explanatory variable in equation (1) that would specify the short-run, dynamic behavior of current variables. It is important to note that the term "equilibrium" as used here does not have anything to do with market clearing or the equality of actual and desired quantities. Rather, it refers to any long-run relationship that may exist among the

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[6] A threshold model incorporates a dummy variable that has been chosen providing the best in-sample fit and avoids the danger of *ex post* selection (see Chan (1993)). A Fourier approximation model is $y_t = \alpha(t) + \varepsilon_t$, where $\alpha(t) = A_0 + \sum_{k=1}^{s}\left(A_k \sin\frac{2\pi k}{T}\cdot t + B_k \cos\frac{2\pi k}{T}\cdot t\right)$. s is the number of frequencies in the process $\alpha(t)$ (Enders and Hoover (2003)).

nonstationary variables, e.g., when $\eta_{t-1} = 0$ in equation (4) below. Engle and Granger (1987) state that this could be defined as any causal, behavioral, or reduced-form relationship among commonly trending variables (Enders (2004)).

Since the effect economic growth has on poverty depends on the extent of inequality, inequality must be controlled for in any poverty function (Adams (2002)). Our specification including the disequilibrium term and Gini coefficient is,

$$\Delta P_t = \gamma + \alpha\eta_{t-1} + \sum_{i=1}^{p} \beta_{Pi}\Delta P_{t-i} + \sum_{i=1}^{p} \beta_{GINIi}\Delta GINI_{t-i} + \sum_{i=1}^{p} \beta_{Ui}\Delta U_{t-i} + \sum_{i=1}^{p} \beta_{TRi}\Delta TR_{t-i} + \sum_{i=1}^{p} \beta_{FEMi}\Delta FEM_{t-i}$$
$$+ \sum_{i=1}^{p} \beta_i \Delta GDP_{t-i} + \sum_{j=1}^{4}\sum_{i=1}^{p} \theta_{i,j}\Delta GDP_{t-i} \cdot D_j + \varepsilon_t, \qquad (2)$$

$$\begin{aligned} D_1 &= 1 \quad \text{when } 1962 \le t \le 1969 \\ &= 0 \quad \text{otherwise,} \\ D_2 &= 1 \quad \text{when } 1971 \le t \le 1973,\ 1976 \le t \le 1979 \\ &= 0 \quad \text{otherwise,} \\ D_3 &= 1 \quad \text{when } 1983 \le t \le 1989 \\ &= 0 \quad \text{otherwise,} \\ D_4 &= 1 \quad \text{when } 1992 \le t \le 1999 \\ &= 0 \quad \text{otherwise,} \end{aligned} \qquad (3)$$

$$\eta_{t-1} = \psi_0 + \psi_1 P_{t-1} + \psi_2 GINI_{t-1} + \psi_3 GDP_{t-1} + \psi_4 U_{t-1} + \psi_5 TR_{t-1} + \psi_6 FEM_{t-1}. \qquad (4)$$

All of the numeric variables are in natural logs which mean that all coefficients may be interpreted as elasticities. The descriptive statistics for the first difference of each variable may be found in Appendices I and II, respectively.[7] Note that in addition to series of lagged values for each variable, equation (2) contains η_{t-1} which represents the long-run relationship among the variables (equation (4)). We incorporated more

[7] The variables were downloaded from http:www.haverselect.com, a fee based on-line service.

expansionary periods than has been utilized in previous studies. The four expansionary periods are the 1960s, 1970s, 1980s, and 1990s (as denoted in (3) above).[8] Since we would expect β_i, the economic growth coefficient, to be negative, if $\theta_{ij} < 0$ then economic growth has reduced poverty more during the *jth* expansionary period. Moreover, if $\theta_4 < \theta_3 < \theta_2 < \theta_1$, then the impact of growth on poverty during economic expansions has diminished over time—or in the vernacular of Thornton, et. al. (1978), "trickle down has petered out." Finally, the specification of equation (2) in an error-correction form assumes that $P_t, GINI_t, GDP_t, U_t, TR_t,$ and FEM_t are cointegrated--their time paths are influenced in a stationary way by the extent of any deviation from long-run equilibrium and α_i is the speed of adjustment parameter.

Estimation and Empirical Results

All of the variables should be integrated of order one ($I(1)$) or have a unit root in order for the specification of equation (2) to be valid. The results of performing the Phillips-Perron unit root tests (Phillips and Perron (1988)) on each explanatory variable are presented in Table I.[9] The null hypothesis of a unit root could not be rejected in each case.

[Insert Table I Here]

Since there is no reason to posit *a priori* that there is asymmetric adjustment in the poverty cointegrating relationship (equation (4)), we use the maximum eigenvalue test developed by Johansen (1996) to test for cointegration. The results are found in

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[8] These periods have been identified as expansions by the National Bureau of Economic Research (http://www.haverselect.com).

[9] The Phillips-Perron tests are used because of the less stringent assumptions regarding the distribution of the error terms in the unit root model specifications.

Table II and only one cointegrating relationship is indicated at $\alpha = .01$. Thus, the variables poverty, gross domestic product, male unemployment rate, federal, state, and local transfer payments per capita, and the number of female-headed households all appear to be cointegrated.

[Insert Table II Here]

Normalizing on P_t, equation (4) may be expressed as:

$$P_t = \lambda_0 + \lambda_1 GINI_t + \lambda_2 GDP_t + \lambda_3 U_t + \lambda_4 TR_t + \lambda_5 FEM_t + \vartheta_t . \qquad (5)$$

Equation (5) was estimated using the Phillips and Hansen (1990) method which amounts to including a differenced term at lag one, zero, and a differenced term of lead one for each explanatory variable.[10] The parameter estimates of the relevant terms are,[11]

Explanatory Variables	Coefficient	Std. Error	z-Statistic	Prob.
GINI	2.927412	0.319084	9.174418	0.0000
GDP	-1.941203	0.200955	-9.659897	0.0000
U	0.098650	0.059252	1.664917	0.0959
TR	-0.185489	0.080803	-2.295575	0.0217
FEM	0.318773	0.162373	1.963214	0.0500
Intercept	17.98609	2.276548	7.900596	0.0000
Sample (adjusted):	**1959-1998**			
Included observations:	**40**			
R-squared	0.923496			
Adjusted R-squared	0.824492	**F-statistic**	9.327817	
S.E. of regression	0.080436	**Prob(F)**	0.000010	

10 This method controls for serial correlation and endogeneity.

11 We did not include the coefficients of all the lead and lag terms for reasons of brevity. Complete results are available upon request from the authors.

All of the explanatory variables are statistically significant at $\alpha = .05$ or $\alpha = .01$, except the male unemployment rate.[12] Thus, unemployment has no impact on poverty when controlling for GDP--it appears that there is a component of male unemployment that is heavily influenced by economic activity. We reached this conclusion because if GDP is omitted and the above equation re-estimated, the coefficient on the male unemployment rate becomes statistically significant at $\alpha = .01$.[13]

From a hierarchical perspective, the level of inequality (GINI), and GDP appear to be the most dominant in explaining long-term movements in the poverty rate. It is important to note that the above results may or may not hold in the short-run—that is, for year-to- year changes (or first differences) in these variables.

The estimation results of the error-correction model (equation (2), above) may be found in Table III. It appears that "trickle down has ***not*** petered out (Thornton, et. al. (1978))." All of the coefficients on the last four variables at lag two, which represent the interactions between the expansion period dummies and the growth in GDP, are statistically significant. In addition, the percentage change in ΔP_t (change in poverty) as a result of a one percent increase in economic growth was for each period,[14]

1960s:	-3.6650 percent,
1970s:	-1.7571 percent,
1980s:	-2.3809 percent,
1990s:	-3.7481 percent.

Our results indicate that contrary to previous studies, the 1980s expansion was significant and the effect of economic growth on poverty has not diminished over

[12] We must note that the coefficient on male unemployment is significant at $\alpha = .05$.

[13] Results are available from the authors upon request.

[14] The elasticities are computed by adding -1.1469, the lag two coefficient of economic growth, to each interaction coefficient for each expansionary period.

time. Indeed, the percentage change in ΔP_t for the 1990s expansion (-3.7481) is greater than the elasticity coefficient for the 1960s expansion (-3.665).

[Insert Table III Here]

As previously mentioned, while the male unemployment rate does not affect poverty in the long-run, changes in this rate do influence changes in poverty in the error-correction equation. The coefficient on the lag one and lag two change in the male unemployment rate is positive and statistically significant at the .05 level. However, while statistically significant, it has a numerically negligible influence on poverty at both lags. At lag one, changes in male unemployment has the least impact on changes in poverty as compared to all of the explanatory variables--the elasticity coefficient is .1273 percent. This result is consistent to what we have mentioned previously concerning the importance of the cyclical component—the effect of changes in the male unemployment rate on changes in poverty is small when controlling for economic growth. It is also important to note that short run empirical result runs counter to what has been replicated in past studies, which have found unemployment to have a rather large affect on poverty rates.[15] According to Thornton, et al. (1978), unemployment is included in the model to control for the "last hired-first fired" syndrome. Our results imply that policies improving economic growth may deliver a larger "bang for the buck" as far as changes in poverty are concerned, than labor market policies dealing directly with male unemployment rates, *ceteris paribus*.

[15] See Anderson (1964), Thornton, et al. (1978), Hirsh (1980), and Formby et al. (2001).

The explanatory variable, change in the number of female-headed households, does not appear to significantly affect changes in poverty in the short run. As mentioned previously, this variable is one of the factors that explain long-term movements in poverty. This result that favors the long-term is not surprising, since the rise in the number of single-parent households over the past 50 years has been a dominant and controversial factor in the composition of poverty rates. One reason given for this increase has to do with the corresponding reduction in poverty among other demographic groups in the economy over the long-term. Increases in Social Security payments have reduced the incidence of poverty among the elderly and the Supplemental Social Security program, introduced in 1973, has reduced poverty among the disabled (Wentworth and Pattison (2002)).

State, local, and federal transfer payments do not only influence poverty over the long-term, but changes in transfers are also found to have a statistically significant and negative affect on changes in poverty in the short-term (at lag one). Proponents of expanding state, local, and/or federal services might take solace in this result, since it appears that there is room, at least on a year-to-year basis and over the long-run, for poverty reduction through increased transfers. But the magnitude of the elasticity coefficient on the lag one transfers variable in Table III is among the smallest of the explanatory variables—meaning that its impact on changes in poverty in the short run is comparatively slight. Regardless of the short-term results, we show that poverty is affected by transfer payments over the long-term--the 40-year period covered by this study.

Conclusion

There are problems pertaining to the poverty measurement variable. We use the standard measure of poverty which is formally defined in the United States in absolute terms and is measured by the number of persons with equivalence scale adjusted incomes below the "Orshansky" poverty line (Orshansky (1965.1, 1965.2). It has been acknowledged, ever since the influential contribution of Sen (1976), that headcount measures of poverty are problematic because there are other aspects of the income distribution that are ignored. For example, if only the headcount matters, income could be redistributed from the poorest of the poor to families slightly below the poverty line and the official poverty measure would decrease. Sen (1976) demonstrates that when the head count ratio and average income shortfall (poverty gap) of the poor are both constant, a rise in income inequality among the poor necessarily increases the economic deprivation among the poor. In their study, Formby, et. al. (2001) utilize the *Sen index* which is sensitive to headcount poverty, the income shortfall of the poor, and the distribution of income among the poor (Formby, et. al. (2001)). However, use of this superior measure did not yield any substantive differences in their econometric results, e.g., the differences in magnitudes and statistical significance of the coefficients between the equation using absolute poverty (headcount measure) as the dependent variable and the equation using the Sen measure were virtually the same.[16] Thus, there is evidence that use of the absolute poverty measure should not invalidate our results.

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[16] See Formby, J.P., G.A. Hoover, and H. Kim (2001), *Economic Growth in the United States:Comparisons of Estimates Based Upon Official Poverty Statistics and Sen's Index of Poverty*, Working Paper, University of Alabama, Table II, Model 5 and 6.

The "trickle-down" effect of economic growth is an important issue in policy debate. Anderson (1964) hypothesized that poverty in America would become less responsive to economic growth and new policies would be needed if poverty were to be reduced. He believed that a considerable proportion of the poor were made up of children, the elderly, and the disabled who were incapable of full-time work. These groups were simply not affected by the poverty-reducing effects of economic growth. Blank and Card (1993) have shown that U.S. government poverty statistics have become less sensitive to economic growth across time. As a consequence, growth is believed to have become less effective as a poverty-fighting tool than it was in the 1960s (Formby, et. al. (2001)).

An alternative hypothesis is that an increasing rate of economic growth has an even greater influence on the reduction of poverty. The implication here is that some workers may not be hired under normal growth conditions, but may find increased employment opportunities during periods of high and sustained economic growth. We have found that increases in economic growth are indeed significantly related to reductions in the poverty rate for all families, *ceteris paribus*. In addition, by using an appropriately specified error-correction model, we have shown that economic growth has had a pronounced effect on poverty during the economic expansions of the 1960s, 1970s, 1980s, and 1990s. This is in contrast to some previous studies that have posited the effect of economic growth on changes in poverty to have either declined or remained unchanged over time, e.g., the most recent being the aforementioned analysis by Formby, et al. (2001). Our results also do not support the contentions of previous analyses that the effect of substantial economic growth on changes in poverty has moderated during the

1980s. During this period, the effect of increased GDP growth on changes in poverty was 1.23 percent higher than during other expansionary periods.

In our approach, we have formulated a model that is more theoretically and statistically valid than what has been used in prior studies. In previous analyses of the growth-poverty relationship, there has never been mention of a "long run" relationship among the variables of interest--nor has the time series properties of the relevant variables ever been analyzed.

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Appendix I

Data Sources

Variable	Years
P - Poverty Rate, All Families	1959 – 1999
GDP – Gross Domestic Product (Chained) (Bils. 92 $)	1946 – 1999
GINI – Gini Coefficient	1947 – 1999
U – Male Unemployment Rate	1948 – 1999
TR – Federal, State and Local Transfer Payments Per Capita (92 $)	1946 – 1999
FEM – Number of Female Headed Households (000s)	1947 – 1999

Appendix II

Descriptive Statistics

First Difference of Logged Variables

Expansionary Period: 1962-1969

	Mean	Median	Std. Deviation	Minimum	Maximum
P	-0.0742	-0.0652	0.0453	-0.1629	-0.0260
GDP	0.0480	0.0515	0.0148	0.0247	0.0636
GINI	-0.0086	-0.0084	0.0190	-0.0326	0.0255
TR	0.1062	0.0999	0.0717	0.0210	0.2155
U	-0.1033	-0.1010	0.0830	-0.2060	0.0016
FEM	0.0335	0.0351	0.0173	0.0087	0.0605

Expansionary Period: 1971-1973, 1976-1979

	Mean	Median	Std. Deviation	Minimum	Maximum
P	-0.0253	-0.0174	0.0312	-0.0696	0.0260
GDP	0.0466	0.0529	0.0105	0.0313	0.0561
GINI	0.0044	0.0055	0.0074	-0.0084	0.0139
TR	0.1575	0.1343	0.0391	0.1178	0.2074
U	-0.0708	-0.1127	0.1275	-0.1775	0.1898
FEM	0.0451	0.0407	0.0096	0.0351	0.0588

Expansionary Period: 1983-1989

	Mean	Median	Std. Deviation	Minimum	Maximum
P	-0.0227	-0.0282	0.0205	-0.0541	0.0132
GDP	0.0418	0.0378	0.0130	0.0334	0.0701
GINI	0.0077	0.0052	0.0055	0.0025	0.0155
TR	0.0841	0.0823	0.0240	0.0526	0.1257
U	-0.0927	-0.0554	0.0978	-0.2880	-0.0042
FEM	0.0179	0.0151	0.0149	0.0012	0.0437

Expansionary Period: 1992-1999

	Mean	Median	Std. Deviation	Minimum	Maximum
P	-0.0221	-0.0351	0.0370	-0.0651	0.0414
GDP	0.0356	0.0373	0.0072	0.0262	0.0434
GINI	0.0097	0.0058	0.0225	-0.0118	0.0600
TR	0.0648	0.0628	0.0324	0.0238	0.1328
U	-0.0694	-0.0957	0.0745	-0.1532	0.0973
FEM	0.0157	0.0205	0.0097	0.0013	0.0264

Table I

Phillips-Perron Unit Root Tests

Variable	PP Test Statistic	P Value
Log of Poverty Rate	-2.33	.1686
Log of Gini	.36	.9784
Log of Real Gross Domestic Product	-1.83	.3595
Log of Male Unemployment Rate	-2.09	.2492
Log of Real State, Local, and Federal Transfer Payments Per Capita	- .83	.7993
Log of Number of Female-Headed Households	-2.86	.0591

Table II

Johansen Maximum Eigenvalue Test

$$H_0 : r$$
$$H_1 : r+1$$

Hypothesized No. of CE(s)	Eigenvalue	Max-Eigen Statistic	5 Percent Critical Value	1 Percent Critical Value
**None ** **	0.8765	79.4830	39.37	45.10
At most 1 *	0.6185	36.6222	33.46	38.77
At most 2 *	0.5444	29.8757	27.07	32.24
At most 3	0.3665	17.3494	20.97	25.52
At most 4	0.3044	13.7913	14.07	18.63
At most 5	0.0243	0.9360	3.76	6.65

Max-eigenvalue test indicates 1 cointegrating equation(s) at the 1% level

***(**) denotes rejection of hypothesis at 5 % (1 %) level**

Table III

Error Correction Model

Estimation Results

Dependent Variable: ΔP_t		**Adjusted R-Squared:**	**.588972**
Sample (adjusted): **1962 1999**		**F-Statistic:**	**3.209090**
Included observations: 38 after adjustments		**Prob (F-Statistic):**	**.016241**

	Coefficient	**Std. Error**	**z-Statistic**	**Prob.**
$\hat{\eta}_{t-1}$	-0.099745	0.060409	-1.651154	0.1000
ΔP_{t-1}	-0.001588	0.000961	-1.651189	0.1000
ΔP_{t-2}	0.610246	0.118989	5.128598	0.0000
$\Delta GINI_{t-1}$	0.417283	0.173719	2.402055	0.0163
$\Delta GINI_{t-2}$	1.095583	0.133190	8.225708	0.0000
ΔGDP_{t-1}	-1.012603	0.612099	-1.654311	0.1000
ΔGDP_{t-2}	-1.146941	0.470558	2.437409	0.0148
ΔU_{t-1}	0.127257	0.045521	2.795586	0.0052
ΔU_{t-2}	0.112684	0.051365	2.193801	0.0282
ΔTR_{t-1}	-0.196023	0.095130	-2.060583	0.0393
ΔTR_{t-2}	-0.030175	0.105563	-0.285846	0.7750
ΔFEM_{t-1}	0.710623	0.537886	-1.321141	0.1868
ΔFEM_{t-2}	0.066802	0.189468	-0.352580	0.7244
D60S* ΔGDP_{t-1}	0.252055	0.450817	0.559107	0.5761
D60S* ΔGDP_{t-2}	-2.518015	0.476337	-5.286203	0.0000
D70S* ΔGDP_{t-1}	-0.081553	0.516571	-0.157873	0.8746
D70S* ΔGDP_{t-2}	-0.610165	0.131481	-4.640709	0.0000
D80S* ΔGDP_{t-1}	-0.481152	0.362143	-1.328625	0.1840
D80S* ΔGDP_{t-2}	-1.233995	0.356405	-3.462340	0.0005
D90S* ΔGDP_{t-1}	0.415301	0.447151	0.928772	0.3530
D90S* ΔGDP_{t-2}	-2.601196	0.512960	-5.070953	0.0000
Intercept	0.040038	0.018964	2.111203	0.0348