STRUCTURAL UNEMPLOYMENT, CYCLICAL UNEMPLOYMENT, AND INCOME INEQUALITY

H. Naci Mocan*

Abstract—This is the first study that decomposes unemployment into its structural and cyclical components and investigates their impact on income distribution, controlling for the influence of inflation. Increases in structural unemployment have a substantial aggravating impact on income inequality. Inflation has a progressive impact, which is due to the unexpected component. The study demonstrates that previous work failed to take into account the stochastic trend behavior of the variables. Consequently, specifications used by previous research cannot predict the behavior of income shares after 1983, whereas the specification used by this paper generates accurate forecasts. The results also indicate that a sustained GNP growth is not necessarily associated with an improvement in income inequality, because sustained GNP growth can coexist with increased structural unemployment.

I. Introduction

THE relationship between macroeconomic conditions and income distribution has long attracted the attention of economists (Schultz, 1969; Metcalf, 1969; Beach, 1977; Blinder & Esaki, 1978; Buse, 1982; Blank & Blinder, 1986; Nolan, 1989; Blejer & Guerrero, 1990; Bjorklund, 1991; Silber & Zilberfarb, 1994; Jäntti, 1994). Although researchers have differed in the exact specification of their models, the general methodology has involved regressing income shares or the Gini coefficient on aggregate macroeconomic performance indicators: typically the unemployment rate and inflation. The consensus has been that income inequality is countercyclical in behavior, i.e., increases in unemployment worsen the relative position of low-income groups. On the other hand, contrary to the popular belief that ''the poor [are] the chief sufferers of inflation,'' [1] the empirical evidence on the impact of inflation is not conclusive. The progressive impact of inflation was reported by Metcalf (1969), Blinder and Esaki (1978), Blank and Blinder (1986), Jäntti (1994), and Bishop et al. (1994) who used data from the United States, and Flückiger and Zarin-Nejadan (1994) with data from Switzerland. Blejer and Guerrero (1990) and Silber and Zilberfab (1994) uncovered a regressive impact in the Philippines and Israel, respectively, whereas Buse (1982) found no relationship between inflation and income inequality in Canada.

Although the cyclical effect of unemployment on income inequality has intuitive appeal, the behavior of income inequality in the United States over the past two decades suggests the importance of more-permanent components. For example, the share of the lowest quintile of family income distribution in the United States declined from 5.5% in 1970 to 4.2% in 1994. On the other hand, the share of the highest quintile rose from 41% to 47% during the same period. To maintain that there exists a trend in income inequality that works in favor of the rich is not particularly informative in and of itself unless the determinants of that trend are explored.

This paper uses the same framework employed by earlier studies to investigate the influence of macroeconomic conditions on income inequality. However, it differs from them in two important ways. The first is through the treatment of the unemployment rate. Previous studies investigated the relationship between inequality and business cycles by including the unemployment rate (or employment) as an explanatory variable; this paper decomposes the unemployment rate into its structural and cyclical components. This is important because it allows a test as to whether changes in short-term (cyclical) or more-permanent (structural) unemployment have dissimilar influences on changes in income inequality. The second improvement pertains to the treatment of the trends. Previous studies attempted to control for secular movements in inequality by including time trends in the regression equations (e.g., Jäntti, 1994; Beach & Slotsve, 1994; Blejer & Guerrero, 1990; Buse, 1982; Blinder & Esaki, 1978; Schultz, 1969). However, there exists an immense body of time-series literature that demonstrates that including time trends into regression analysis may generate misleading results, if the variables in the analysis contain stochastic trends. This paper suggests that using trend terms in inequality regressions is indeed inappropriate in the light of new developments in time-series econometrics.

Section II describes the background and the unit-root tests. Section III introduces structural unemployment into the analysis and decomposes unemployment into its structural and cyclical components. Section IV reports the estimated income inequality regressions and the forecasts of the competing models. Section V is the conclusion.

II. Background and Levels versus Differences

Previous studies investigated the impact of macro conditions on income inequality through estimating regressions of the following form:

$$S_{i,t} = \alpha_i + \beta_i U_t + \gamma_i \pi_t + \delta_i t + \epsilon_{i,t}, \quad i = 1, 2, \ldots, 5 \qquad (1)$$

where $S_{i,t}$ is the share of the *i*th quintile in the distribution of income among U.S. families in year *t*, U is the unemploy-

Received for publication April 27, 1995. Revision accepted for publication January 20, 1998.

* University of Colorado at Denver and National Bureau of Economic Research, New York, NY.

An earlier version of this paper was presented at the 1995 American Economic Association Meetings in Washington, D.C. I thank Jeffrey Zax, Dan Hamermesh, Gary Solon, Madeline Mocan, and Daniel I. Rees for helpful suggestions; Chris Underwood, Roger Johnson and Erdal Tekin for research assistance; and Leyla Mocan for inspiration. Anonymous referees provided very valuable comments.

[1] Arthur Burns, as cited by Palmer (1973) and Blinder and Esaki (1978).

The Review of Economics and Statistics, February 1999, 81(1): 122–134

TABLE 1.—DICKEY-FULLER TESTS[a]

	Levels of the Variables						
Model	Lowest Quintile	Second Quintile	Middle Quintile	Fourth Quintile	Highest Quintile	Inflation	Unemployment Rate
2A	−0.894	1.071	1.622	−2.664	1.526	−2.171	−2.550
2B	−1.019	−1.475	−0.638	−1.809	−0.047	−1.761	−3.542
				Differences of the Variables			
Model	Lowest Quintile	Second Quintile	Middle Quintile	Fourth Quintile	Highest Quintile	Inflation	Unemployment Rate
2A	−5.509	−4.606	−4.088	−6.442	−4.136	−7.064	−5.988
2B	−6.146	−5.503	−5.505	−7.030	−5.703	−7.125	−5.911

Notes: [a] The entries are the calculated test statistics of γ for the corresponding model. MacKinnon critical values for 46 observations are −3.578, −2.926 and −2.600 at 1%, 5% and 10%, respectively for Model 2A, and −4.168, −3.509 and −3.184 for Model 2B.

ment rate, π stands for the rate of inflation, t is a linear time trend, and ϵ is a white-noise disturbance with usual properties. Researchers estimated alternative versions of the model which included a quadratic unemployment term or a lagged dependent variable as additional regressors. They also employed different measures of unemployment and inflation. An earlier version of this paper replicated the results of Blinder and Esaki (1978), Blank and Blinder (1986), and Jäntti (1994). The analyses presented in that paper (Mocan, 1995) demonstrated that, within the framework of earlier research, increases in unemployment worsen income inequality, and increases in inflation help to reduce it. These results were insensitive to the time period under investigation, the exact measure of independent variables, the functional form of the model (inclusion of the quadratic terms or lagged dependent variables), or the estimation procedure (OLS or GLS). These results are available upon request.

It is well known that the usual techniques of regression analysis can result in highly misleading conclusions when variables contain stochastic trends (Stock & Watson, 1988; Nelson & Kang, 1981; Granger & Newbold, 1974). In particular, if the dependent variable and at least one independent variable contain stochastic trends, and if they are not cointegrated, the regression results are spurious (Phillips, 1986; Granger & Newbold, 1974). To identify the correct specification of the model depicted by equation (1), an investigation of the presence of stochastic trends in the variables is needed. To this end, augmented Dickey-Fuller tests described by equations (2A) and (2B) are applied to the variables of the model. They are the shares of family income received by income quintiles, the overall unemployment rate, and inflation based on the implicit GNP deflator.[2] The data span the years 1948 to 1994.

$$\Delta X_t = \mu_1 + \gamma X_{t-1} + \psi_1 \Delta X_{t-1} + \epsilon_t, \tag{2A}$$

$$\Delta X_t = \mu_1 + \mu_2 t + \gamma X_{t-1} + \psi_1 \Delta X_{t-1} + \epsilon_t. \tag{2B}$$

X_t is the series under investigation, and Δ stands for the first-difference. If $\gamma = 0$, then X_t contains a unit root and is therefore an I(1) process, governed by a stochastic trend. Since the estimated γ does not have the usual asymptotic distribution, the values tabulated by MacKinnon (1991) are used; these values are more accurate than the ones originally tabulated by Fuller (1976) and Dickey and Fuller (1981). The results are reported in the first panel of table 1. The calculated test statistics for income shares are always larger than the critical values with the exception of the fourth quintile, which is significant at the 10% level in Model 2A. Thus, based on Dickey-Fuller tests, the hypothesis of a unit root (stochastic trend) cannot be rejected for income shares, which is consistent with the results reported by Hayes et al. (1990). The same is true for the inflation rate. For the unemployment rate, the issue is less clear. While the result from Model 2A suggests the presence of a unit root, Model 2B rejects the hypothesis of a unit root at the 5% level.

There exists a consensus among macroeconomists that the unemployment rate does not have a unit root. As a result, some recent macroeconomic research does not even test for the presence of stochastic trends in the unemployment rate (e.g., Perron, 1989, p. 1363; Blanchard & Quah, 1989, p. 660). On the other hand, there are cases in which researchers were unable to reject the hypothesis of a unit root in the unemployment rate (e.g., Joyce & Mocan, 1993; Macunovich & Easterlin, 1988). To analyze the issue further, following Stock (1991), 95% confidence intervals and median-unbiased point estimates for the largest autoregressive roots are obtained, which are presented in table 2. The first row for each variable reports the confidence interval and the median estimate for Model 2A; the second row pertains to Model 2B. In all cases, the confidence intervals contain the unit

[2] Unemployment and GNP deflator data are obtained from Citibase. Income shares are obtained from the Current Population Reports, Consumer Income Series. There are well known measurement problems with the income share data as reported by Blinder and Esaki (1978). The changing family structure over time introduces another difficulty. It is nevertheless important to use the family income quintiles to make the results of this paper comparable to those of earlier studies that used the same exact measure. Furthermore there is evidence indicating that the contribution of changing family structure on inequality and poverty is modest (Cutler and Katz 1991).

TABLE 2.—CONFIDENCE INTERVALS AND MEDIAN-UNBIASED ESTIMATES FOR THE LARGEST AUTOREGRESSIVE ROOT[a]

Variable	95% Interval	Median-Unbiased Estimate
Lowest Quintile	(0.90, 1.10)	1.03
	(0.93, 1.11)	1.05
Second Quintile	(0.95, 1.10)	1.04
	(0.84, 1.11)	1.05
Middle Quintile	(0.88, 1.10)	1.03
	(1.03, 1.11)	1.06
Fourth Quintile	(0.49, 1.04)	0.75
	(0.77, 1.10)	1.04
Highest Quintile	(1.03, 1.12)	1.05
	(1.03, 1.12)	1.06
Inflation	(0.62, 1.06)	0.86
	(0.77, 1.11)	1.04
Unemployment Rate	(0.51, 1.04)	0.77
	(0.24, 1.05)	0.58

Note: [a] The first row for each variable reports the confidence interval and the median estimate for the model presented in equation (2A), the second row pertains to equation (2B).

root, but they are much wider for the unemployment rate. Furthermore, the unit root is much closer to the upper bound of the confidence interval, and the median-unbiased estimates are less than 1 for the unemployment rate in both models. Given this supporting evidence provided in table 2, the hypothesis of a unit root is rejected for the unemployment rate, but it is maintained for income shares and inflation.[3] The second panel of table 1 displays the results of the unit root tests for the first-differences of the variables. The hypothesis of a unit root is rejected in all cases, indicating that the variables are not I(2) processes.

Given the evidence provided by Hendry and Neale (1991) that regime shifts can mimic unit roots in autoregressive time series,[4] and given that Perron (1989) rejects the hypothesis of a unit root for most U.S. macroeconomic time series in the model that includes a break in the trend, it is important to investigate whether income shares and inflation are indeed governed by stochastic trends or whether breaks in their underlying trends are responsible for the appearance of the unit root. Following Zivot and Andrews (1992)—who extended Perron's test where the breakpoint in estimated, rather than fixed—and Raj and Slottje (1994)—who applied the test to several inequality measures—the following regression is estimated:

$$X_t = \mu + \beta t + \theta DU_t(\lambda) + \gamma DT_t(\lambda) + \rho X_{t-1} + \sum_{j=1}^{k} c_j \Delta X_{t-j} + \epsilon_t, \tag{3}$$

where DU_t is a level-shift dummy; $DU_t = 0$ if $t \le T_B$, and $DU_t = 1$ otherwise; DT_t is a slope-shift dummy; $DT_t = t - T_B$ if $t > T_B$ and $DT_t = 0$ otherwise; T_B is the unknown year in which the structural change took place; and $\lambda = T_B/T$.

[3] Stock (1991) reports similar 90% confidence intervals for the unemployment rate using the annual Nelson-Plosser data spanning 1890–1970. In our case, the 90% confidence interval for the unemployment rate is (0.30, 0.94).

[4] I thank an anonymous referee for this insight.

This model allows for segmented trends with level and slope shifts at the breakpoint T_B. The breakpoint is unknown and determined endogenously, so that it gives the least-favorable result to the null hypothesis of a unit root. That is, $\lambda = T_B/T$ is chosen to minimize the one-sided t-statistic for testing $\rho = 1$ (Zivot & Andrews, 1992). If $\hat{\lambda}_{inf}$ stands for a minimizing value for the model, then $t_\rho[\hat{\lambda}_{inf}] = \inf t_\rho(\lambda)$, where λ is an element of a closed subset Ω, which consists of values created by $T_B = 1951, 1952, \ldots, 1994$. The lag length k is determined using the same selection procedure used by Perron (1989), Zivot and Andrews (1992) and Raj and Slottje (1994). The results are reported in table 3. The critical values tabulated by Zivot and Andrews (1992) are -5.57 at the 1% level, -5.08 at the 5% level, and -4.82 at the 10% level. In no case can we reject the hypothesis of a unit root, even with the breaks in trends and jumps in levels that occurred in the years presented in column 2 of table 3.

In summary, the analysis described above provides evidence against the unit root for the unemployment rate, but the hypothesis of a unit root cannot be rejected for inflation and income shares. The presence of a unit root implies that the variance of the series is a function of time, and the variance would infinitely increase over time. This creates a conceptual difficulty for income shares, because they are bounded series. Given that it is not possible to determine the exact structure of the underlying data-generating mechanism with a finite sample, the evidence for a unit root, and therefore the need to employ the series in first-difference form, can be considered as an approximation. Later in the paper, I show that alternative formulations for income shares, such as deterministic trends and segmented trends, provide poorer predictive performance of the model.

Although income shares and inflation are governed by stochastic trends, they are cointegrated if there exists a linear combination of them which is stationary with zero mean. Two versions of the cointegration test are performed by running the regressions

$$S_{i,t} = \alpha + \beta\pi_t + \epsilon_{i,t}, \text{ and} \tag{4A}$$

$$S_{i,t} = \alpha + \beta\pi_t + \delta t + \epsilon_{i,t}, \; i = 1, \ldots, 5. \tag{4B}$$

and investigating whether the residuals $\hat{\epsilon}_{i,t}$ contain a unit root. This is done by running the cointegration regression

$$\Delta\hat{\epsilon}_{i,t} = \eta\hat{\epsilon}_{i,t-1} + \omega_{i,t}, \tag{5}$$

where $\omega_{i,t}$ is a white-noise error term. If the estimated η is zero, $\hat{\epsilon}_{i,t}$ has a random-walk behavior, which indicates that the variables (which are governed by stochastic trends) do not share a common trend; i.e., they are not cointegrated. The estimated t-values for η are reported in table 4. Although η based on Model 4A is significant for the fourth quintile at the 5% level, it is not different from zero if the cointegration regression is 4B. Similarly, there is no evidence of cointegration between other income shares and

TABLE 3.—UNIT-ROOT TESTS IN THE PRESENCE OF A STRUCTURAL BREAK[a]

Variable	$T_B = \lambda/T$	k	μ	β	θ	γ	t_ρ
Lowest Quintile	1973	1	3.052	0.031	0.043	−0.077	−4.803
			(4.739)	(4.334)	(0.475)	(−4.730)	
Second Quintile	1965	1	6.082	−0.007	0.260	−0.034	−3.547
			(3.610)	(−0.829)	(2.426)	(−2.377)	
Middle Quintile	1981	1	7.419	−0.005	−0.056	−0.046	−2.865
			(2.900)	(−1.509)	(−0.503)	(−2.657)	
Fourth Quintile	1981	4	16.140	0.008	0.314	−0.051	−3.434
			(3.453)	(1.729)	(2.739)	(−3.779)	
Highest Quintile	1977	1	22.664	−0.015	−0.039	0.190	−2.854
			(2.821)	(−1.022)	(−0.129)	(3.037)	
Inflation	1983	2	−0.221	0.123	−2.673	−0.144	−3.023
			(−0.408)	(2.799)	(−2.702)	(−1.060)	
Unemployment Rate	1982	1	2.084	0.061	1.056	−0.204	−4.306
			(3.050)	(2.900)	(1.433)	(−2.423)	

Notes: [a] The figures in parentheses are the *t*-values. t_ρ is the *t*-statistic for the test of the null hypothesis that $\rho = 1$ in equation (3). The critical values are −5.57 at the 1% level, −5.08 at the 5% level, and −4.82 at the 10% level.

TABLE 4.—COINTEGRATION TESTS[a]

Model	Lowest Quintile	Second Quintile	Middle Quintile	Fourth Quintile	Highest Quintile
4A	−1.219	0.429	0.760	−4.031	0.614
4B	−2.848	−3.505	−3.711	−3.332	−3.540

Notes: [a] The entries are the calculated test statistics for η in the corresponding model. MacKinnon critical values for 46 observations are −3.472 and −4.143 at the 5% and 1% level, respectively, for Model 4A with two variables. The 5% and 1% critical values for Model 4B are −3.978 and −4.680 for two variables, and −4.387 and −5.093 for three variables.

inflation. This implies that there is no long-run equilibrium relation between income inequality and inflation. Income inequality can keep deteriorating or keep improving irrespective of the behavior of inflation; i.e., there is not an attractor to hold them together in the long-run (Engle & Granger, 1991).

These results suggest that the models estimated in income inequality literature, displayed by equation (1), have been estimated inappropriately. Since income shares and the rate of inflation contain stochastic trends as revealed by unit root tests, equation (1)—which regresses the level of income shares on unemployment and a trend term—not only produces a random R^2, but the residuals from this regression exhibit spurious periodicity (Nelson & Kang, 1981). Furthermore, inferences based on usual *t*-statistics are not valid (Stock & Watson, 1988). Thus, the proper specification of equation (1), which takes into account the recent developments in time-series econometrics, should involve regressing the first difference of income shares on the first difference of inflation and should not include a time trend as a regressor.

III. Structural Unemployment

If marginal workers with relatively low skills are the ones who are laid off first during an economic downturn, and if these workers are at the bottom part of the income distribution, temporary increases in unemployment are expected to worsen income inequality. On the other hand, the loss of income due to the transitory unemployment of a family member may be offset by unemployment insurance and welfare benefits, especially given the growing incidence of dual earners within a family. Thus, it may take longer spells of unemployment to have a marked effect on annual family incomes. It is therefore important to investigate the possible differential impacts of short-term and long-term unemployment on income distribution.

According to the standard textbook explanation, unexpected demand and supply disturbances bring about changes in production and unemployment. They also generate changes in the rate of inflation, which would bring unemployment back to its long-run level. Long-run unemployment may evolve over time due to changes in technology, the composition of the labor force, and the institutional characteristics of the labor market. Figure 1 displays the behavior of the U.S. unemployment rate and the rate of inflation for the years 1948 through 1994. As Blanchard and Summers (1988) argue, if long-run (structural) unemployment were constant, an increase in the actual unemployment rate should have been associated with a drop in the inflation rate. This is clearly not the case. For example, between 1969 and 1981, the rate of unemployment rose from 3.5% to 7.6%. The inflation rate rose also during the same period from 5% to

FIGURE 1.—UNEMPLOYMENT AND INFLATION

FIGURE 2.—ACTUAL AND STRUCTURAL UNEMPLOYMENT

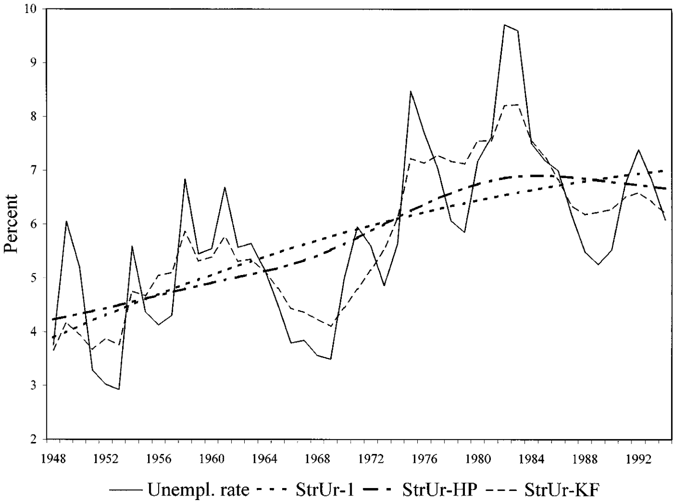
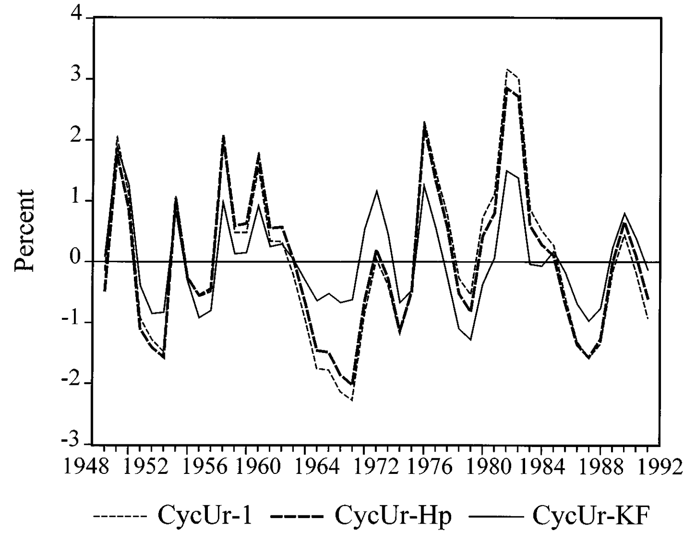


FIGURE 3.—CYCLICAL UNEMPLOYMENT



9.5%. Structural unemployment can change over time due to changes in the demand for labor or due to changes in returns to nonwork. It can also change if the composition of the labor force changes. For example, if the share of teenagers in the labor force rises, this increases the structural rate because teenagers have a higher unemployment rate than the prime-aged workers. An increase in worker mobility between sectors may also generate an increase in structural unemployment (Black, 1982; Lilien, 1982).

In order to investigate whether long-term and short-term unemployment have differential impacts on income inequality, actual unemployment is decomposed into its trend and cyclical components. Because the hypothesis of a unit root is rejected for the unemployment rate, the conventional way to determine the level of structural (long-run) unemployment is to regress the unemployment rate on a constant, and linear and quadratic trend terms. The fitted values represent the long-term (structural) unemployment, whereas the trend deviations illustrate cyclical unemployment. Structural unemployment obtained from this method (StrUr-1) is depicted in figure 2 along with the actual unemployment rate. To assure that the results are not sensitive to the method of decomposition, structural unemployment is also obtained in two alternative ways. First, the Hodrick-Prescott filter (Hodrick & Prescott, 1980) is used to obtain the structural component (Blackburn & Ravn, 1992; Danthine & Girardin, 1989), which provided the path depicted by StrUr-HP in figure 2.[5] Also, following Mocan (1994), Harvey (1985), and Harvey and Todd (1983), the unemployment rate is modeled as governed by a locally linear trend, a cycle and noise component. Application of the Kalman filter allows an estimate of the trend at all points in the sample using all the observations, which is depicted by StrUr-KF in figure 2.[6]

TABLE 5.—CORRELATIONS OF STRUCTURAL AND CYCLICAL UNEMPLOYMENT MEASURES

	Struct. Ur1	Cyclical Ur1	Struct. KF	Cyclical KF	Struct. HP	Cyclical HP
Struct. Ur1	1.000	−0.005	0.763	−0.093	0.975	−0.013
Cyclical Ur1		1.000	0.532	0.798	0.105	0.989
Struct. KF			1.000	0.141	0.841	0.490
Cyclical KF				1.000	−0.066	0.839
Struct. HP					1.000	0.068
Cyclical HP						1.000

Structural unemployment obtained from the Hodrick-Prescott filter and the one obtained from fitting linear and quadratic trends are similar to each other. StrUr-KF is more cyclical than the other two. Structural unemployment obtained from the fitted values of linear and quadratic trends should be considered the standard benchmark case. However, as Table 6 demonstrates, the results obtained from models with other measures of structural unemployment are similar to the ones obtained from the model with standard decomposition (StrUr-1). Figure 2 suggests that structural unemployment was 4% in 1949. It went up to 5.8% in the early 1970s and to 6.8% in the 1990s.

Figure 3 displays the behavior of cyclical unemployment derived from the methods described above. CycUr-1 stands for the cyclical component in the unemployment rate obtained from the model in which structural unemployment is based on linear and quadratic trends. CycUr-HP corresponds to the HP filter, and CycUr-KF is the one that corresponds to the Kalman filter decomposition. CycUr-1 and CycUr-HP are almost identical in location and magnitude, while CycUr-KF is generally smaller in magnitude than the others.

Table 5 presents the zero-order correlations between

[5] Following previous examples (e.g. Blackburn and Ravn 1992), the weight on squared second difference of the growth component, which penalizes acceleration in the trend, is set to be 1600.

[6] Based on the results of table 2, the median-unbiased estimate of the autoregressive coefficient for the model with a trend is imposed on the unemployment rate, which produced a third measure of structural unemployment. The results obtained from this specification were similar to the ones reported in tables 6 and 7.

TABLE 6.—INCOME INEQUALITY REGRESSIONS[a]

Explanatory Variables	Lowest Quintile	Second Quintile	Middle Quintile	Fourth Quintile	Highest Quintile
	Structural Unemployment from Fitted Trend				
Constant	0.210	0.295*	0.393**	0.285	−1.165
	(1.609)	(1.914)	(2.277)	(1.540)	(−2.240)
Structural Unemployment	−0.038	−0.058**	−0.073**	−0.049	0.216**
	(−1.594)	(−1.988)	(−2.180)	(−1.476)	(−2.216)
Cyclical Unemployment	−0.026	−0.041**	−0.014	0.002	0.068
	(−1.516)	(−2.039)	(−0.545)	(0.090)	(1.043)
Δ Inflation	0.051**	0.031**	0.030*	−0.007	−0.107**
	(4.301)	(2.234)	(1.931)	(−0.422)	(−2.372)
R^2	0.45	0.33	0.22	0.06	0.28
Durbin-Watson	2.34	2.26	2.47	2.25	2.55
Q(10)	9.03	3.49	7.03	10.57	7.89
	(0.58)	(0.98)	(0.72)	(0.39)	(0.64)
	Structural Unemployment from the HP Filter				
Explanatory Variables	Lowest Quintile	Second Quintile	Middle Quintile	Fourth Quintile	Highest Quintile
Constant	0.218*	0.240	0.320*	0.227	−0.987**
	(1.730)	(1.605)	(1.887)	(1.254)	(−2.030)
Structural Unemployment	−0.040*	−0.049*	−0.061*	−0.039	0.186**
	(−1.718)	(−1.785)	(−1.937)	(−1.226)	(2.042)
Cyclical Unemployment	−0.024	−0.047**	−0.017	−0.0004	0.077
	(−1.288)	(−2.138)	(−0.692)	(−0.016)	(1.077)
Δ Inflation	0.051**	0.029**	0.028*	−0.009	−0.102**
	(4.337)	(2.089)	(1.759)	(−0.522)	(−2.226)
R^2	0.45	0.32	0.20	0.04	0.27
Durbin-Watson	2.35	2.22	2.40	2.21	2.49
Q(10)	9.30	3.32	6.07	9.85	7.14
	(0.50)	(0.97)	(0.81)	(0.45)	(0.71)
	Structural Unemployment from the Kalman Filter				
Explanatory Variables	Lowest Quintile	Second Quintile	Middle Quintile	Fourth Quintile	Highest Quintile
Constant	0.156	0.170	0.193	0.126	−0.607
	(1.660)	(1.545)	(1.507)	(0.925)	(−1.654)
Structural Unemployment	−0.029*	−0.037	−0.039*	−0.022	0.120*
	(−1.680)	(−1.639)	(−1.703)	(−0.931)	(1.765)
Cyclical Unemployment	−0.036	−0.082**	−0.027	−0.003	0.135
	(−1.214)	(−2.367)	(−0.667)	(−0.070)	(1.174)
Δ Inflation	0.049**	0.024*	0.024	−0.011	−0.087*
	(4.012)	(1.669)	(1.469)	(−0.647)	(−1.838)
R^2	0.45	0.34	0.18	0.03	0.25
Durbin-Watson	2.31	2.20	2.33	2.17	2.40
Q(10)	8.54	4.02	5.29	9.04	6.30
	(0.58)	(0.95)	(0.87)	(0.53)	(0.79)

Notes: [a] The entries in parentheses under the estimated coefficients are the calculated *t*-statistics. Δ stands for the first-difference. * indicates significance at the 10% level or better by a two-tailed test. ** indicates significance at the 5% level or better by a two-tailed test. Q(10) is the Ljung-Box statistics for the null hypothesis that the first ten autocorrelations of the estimated residuals are zero. The numbers in parentheses under the Q-statistics are the marginal significance levels.

structural and cyclical unemployment measures. StrUr-1 and StrUr-HP are very highly correlated with a correlation coefficient of 0.975. Similarly, the correlation coefficient between StrUr-KF and StrUr-HP is 0.841. The correlation between StrUr-1 and StrUr-KF is 0.763, which is less than that of others, but still sizable. The three measures of cyclical unemployment are also highly correlated with each other, with coefficients ranging from 0.798 to 0.989. Table 5 also demonstrates that the cyclical and structural measures of unemployment (that will be employed as separate regressors) are not correlated with each other.

Application of the Dickey-Fuller unit-root tests to structural and cyclical unemployment measures revealed no unit roots. The calculated test statistics were −3.727, −6.558, and −4.032 for CycUr-1, CycUr-KF, and CycUr-HP, respectively, for the model without a trend (2A); and −3.529 and −11.790 for StrUr-KF and StrUr-HP, respectively, for the model with a trend (2B). Because of collinearity, StrUr-1 could not be run with a trend term. In the model without a trend, its calculated test statistics was −43.890.

IV. Estimation Results and Forecasts of Income Shares

Table 6 reports the estimation results of the models in which changes in income shares are regressed on changes in inflation, and the levels of structural and cyclical unemployment. The first panel presents the results when structural unemployment is obtained through fitting linear and quadratic trend terms to actual unemployment (StrUr-1 in figure 2). The second panel is the model where structural unemployment is obtained though the HP filter, and the third panel pertains to the case where structural unemployment is obtained by the Kalman filter (StrUr-KF). The models are estimated by OLS to be comparable to previous studies. The standard errors are corrected following the procedure suggested by Pagan (1984). In all specifications, an increase in structural unemployment is associated with an increase in the share of the highest quintile, but a change in cyclical unemployment has no impact on the income share of this group. The first panel demonstrates that an increase in structural unemployment is associated with reductions in the shares of the second and the middle quintiles. The coefficient of the structural unemployment is negative for the lowest quintile also, but significant at the 12% level by a two-tailed test. In the second panel, the coefficient of structural unemployment is negative and significantly different from zero for the bottom three quintiles. According to the third panel, an increase in structural unemployment reduces the income shares of the bottom three quintiles, where structural unemployment is significant at the 11% level for the second quintile. Thus, table 6 provides evidence indicating that structural unemployment is a significant determinant of income inequality. An increase in structural unemployment is associated with an increase in the income share of the richest twenty percent of the population, and with a decrease in the shares of the bottom three quintiles.

Table 6 also indicates that an increase in cyclical unemployment is associated with a reduction in the income share of the second quintile. An increase in the rate of inflation is associated with an improvement in income inequality as it transfers income from the richest quintile to the poorest sixty percent of the population. This result is consistent with previous research that analyzed data from the United States. In this group are Blank and Blinder (1986), who decomposed inflation into anticipated and unanticipated components. They tested and could not reject the hypothesis that their coefficients were equal. Thus, they combined the two variables into actual inflation, which proved to be positive and significant for the second quintile. Following Blank and Blinder, inflation is decomposed into its expected and unexpected components using an ARIMA (2,1,0) model. One-step-ahead forecasts are employed as anticipated inflation, and the difference between actual and anticipated inflation is the unanticipated inflation.[7] The results, which are presented in table 7, reveal that employing anticipated and unanticipated inflation instead of actual inflation generates a reduction in the statistical significance of the estimated coefficients. The coefficient of anticipated inflation is never different from zero. On the other hand, the coefficient of unanticipated inflation is positive for the bottom three quintiles and negative for the highest quintile in all three models. This implies that it is the unexpected component of inflation that has an equalizing effect on income distribution.[8] This result is consistent with the model of Jovanovic and Ueda (1997), who show that, within a principle-agent framework where price expectations are built into contracts, surprise inflation increases the agent's (labor's) share in output and decreases the principal's (employer's) share.

These results are different from earlier studies that highlighted the substantial negative impact of recessions on income inequality. The results also provide some explanation for the puzzling behavior of income inequality and poverty during the expansion of the 1980s. As Cutler and Katz (1991) summarize, the 1980s witnessed a deterioration in income equality and poverty, despite strong economic expansion demonstrated by the growth in real GNP. They state that ''although the substantial increases in income inequality and poverty between 1979 and 1983 are not surprising given the deep recession of the early 1980s, the continued widening of the income distribution and the sluggish decline in the poverty rate during the macroeconomic expansion of 1983–1989 present a sharp break from the postwar pattern'' (Cutler & Katz, 1991, p. 14). To emphasize this point, they estimate income share equations for the period 1947–1983 and obtain predicted income shares for the period 1947–1989. They show that the model's predictions break down after 1983. More precisely, the forecasts of the model are inaccurate after 1983, if they are based upon the parameters estimated between 1947 and 1983. This is attributed to structural changes regarding the relationship between macroeconomic conditions and income inequality. Put differently, the notion that rising tides lift all boats seems not to be in effect any longer.

I repeat the same exercise and estimate five sets of income share equations between 1948 and 1983 and obtain the forecasts for the period from 1948 to 1994 using the estimated coefficients. The first one of these equations is the one used by earlier studies. It includes inflation, unemployment, unemployment squared, a linear trend term, and a constant as explanatory variables. This is referred to as the ''traditional model'' in the remainder of the paper.[9] Even though the hypothesis of a unit root could not be rejected for income shares, the second model expresses income shares as having been governed by segmented trends. The trend for each income quintile is modeled using equation (3); i.e., the breakpoints are determined based on the results presented in table 3. This formulation gave rise to trends in income shares presented by figures 4A through 4E. This second model,

[7] Application of the Dickey-Fuller tests revealed the presence of a unit-root in the anticipated inflation, but no unit-root in unanticipated inflation. The test statistics were −2.267 and −2.119 for anticipated inflation in Models 2A and 2B, respectively. They were −3.938 and −4.094 for unanticipated inflation.

[8] Estimation of the models with unexpected inflation only provided results similar to the ones reported in table 6.

[9] Inclusion of a dummy variable starting in 1981 or 1983 did not alter the results. Cutler and Katz (1991) used a similar model, which included a lagged dependent variable and real per capita GNP as additional regressors.

TABLE 7.—INCOME INEQUALITY REGRESSIONS WITH INFLATION DECOMPOSITION [a]

Explanatory Variables	Lowest Quintile	Second Quintile	Middle Quintile	Fourth Quintile	Highest Quintile
	Structural Unemployment from Fitted Trend				
Constant	0.158	0.169	0.331*	0.313	−0.995*
	(1.042)	(0.917)	(1.662)	(1.435)	(−1.720)
Structural Unemployment	−0.030	−0.038	−0.063*	−0.053	0.188*
	(−1.054)	(−1.180)	(−1.774)	(−1.408)	(1.792)
Cyclical Unemployment	−0.018	−0.040*	−0.005	0.016	0.033
	(−0.987)	(−1.794)	(−0.196)	(0.620)	(0.472)
Δ Anticipated Inflation	0.015	0.001	−0.015	−0.007	0.017
	(−0.945)	(0.073)	(−0.709)	(−0.318)	(0.286)
Unanticipated Inflation	0.049**	0.033	0.060**	0.029	−0.178**
	(2.784)	(1.542)	(2.561)	(1.143)	(−2.625)
R^2	0.31	0.26	0.29	0.09	0.30
Durbin-Watson	2.21	2.26	2.38	2.19	2.49
Q(10)	9.88	2.30	5.37	11.00	6.55
	(0.45)	(0.99)	(0.87)	(0.36)	(0.77)
	Structural Unemployment from the HP Filter				
Explanatory Variables	Lowest Quintile	Second Quintile	Middle Quintile	Fourth Quintile	Highest Quintile
Constant	0.175	0.126	0.239	0.218	−0.775
	(1.253)	(0.741)	(1.282)	(1.068)	(−1.436)
Structural Unemployment	−0.033	−0.031	−0.049	−0.037	0.151
	(−1.252)	(−1.050)	(−1.496)	(−1.095)	(1.606)
Cyclical Unemployment	−0.015	−0.046*	−0.009	0.013	0.042
	(−0.753)	(−1.912)	(−0.355)	(0.446)	(0.557)
Δ Anticipated Inflation	0.015	0.001	−0.016	−0.008	0.020
	(0.957)	(0.048)	(−0.794)	(−0.388)	(0.346)
Unanticipated Inflation	0.050**	0.031	0.057**	0.026	−0.171**
	(2.845)	(1.460)	(2.405)	(1.019)	(−2.502)
R^2	0.32	0.26	0.27	0.07	0.29
Durbin-Watson	2.23	2.24	2.31	2.14	2.42
Q(10)	10.27	2.16	4.49	9.97	5.67
	(0.42)	(0.99)	(0.92)	(0.44)	(0.84)
	Structural Unemployment from the Kalman Filter				
Explanatory Variables	Lowest Quintile	Second Quintile	Middle Quintile	Fourth Quintile	Highest Quintile
Constant	0.100	0.029	0.080	0.083	−0.282
	(0.913)	(0.235)	(0.549)	(0.519)	(−0.669)
Structural Unemployment	−0.020	−0.015	−0.021	−0.014	0.068
	(−1.090)	(−0.596)	(−0.854)	(−0.559)	(0.945)
Cyclical Unemployment	−0.027	−0.105**	−0.030	0.014	0.124
	(−0.798)	(−2.698)	(−0.659)	(0.279)	(0.945)
Δ Anticipated Inflation	0.007	−0.003	−0.022	−0.007	0.041
	(0.414)	(−0.139)	(−0.972)	(−0.282)	(0.611)
Unanticipated Inflation	0.047**	0.023	0.052**	0.024	−0.153**
	(2.573)	(1.093)	(2.129)	(0.877)	(−2.166)
R^2	0.31	0.32	0.24	0.04	0.27
Durbin-Watson	2.17	2.30	2.25	2.09	2.37
Q(10)	10.05	3.68	4.09	9.48	4.64
	(0.44)	(0.96)	(0.94)	(0.49)	(0.91)

Notes: [a] The entries in parentheses under the estimated coefficients are the calculated *t*-statistics. Δ stands for the first-difference.
* significant at the 10% level or better by a two-tailed test.
** significant at the 5% level or better by a two-tailed test. Q(10) is the Ljung-Box statistics for the null hypothesis that the first ten autocorrelations of the estimated residuals are zero. The numbers in parentheses under the Q-statistics are the marginal significance levels.

called a ‘‘segmented-trends model,’’ employs the unemployment rate, unemployment squared, and the inflation rate as explanatory variables, in addition to the trends depicted in figures 4A through 4E. Thus, the segmented-trends model is an intermediate case between the linear trend model and the model in which income shares are used in first differences. The last three models pertain to equations presented in table 6. They are the models in which income shares and inflation are used in first differences, and the unemployment rate is decomposed into its structural and cyclical components. Model 1 is the one in which StrUr-1 and CycUr-1 are the measures of structural and cyclical unemployment. Model 2 is the one with StrUr-HP and CycUr-HP, and Model 3 is the model with StrUr-KF and CycUr-KF. All models are estimated by OLS. Serial correlation in the errors of the traditional and segmented-trend models are corrected by

FIGURE 4A.—SHARE OF THE LOWEST QUINTILE AND ITS ESTIMATED SEGMENTED TREND

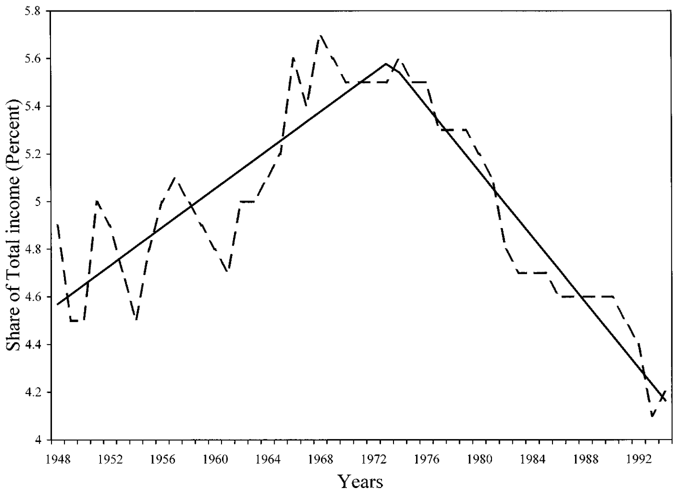


FIGURE 4B.—SHARE OF THE SECOND QUINTILE AND ITS ESTIMATED SEGMENTED TREND

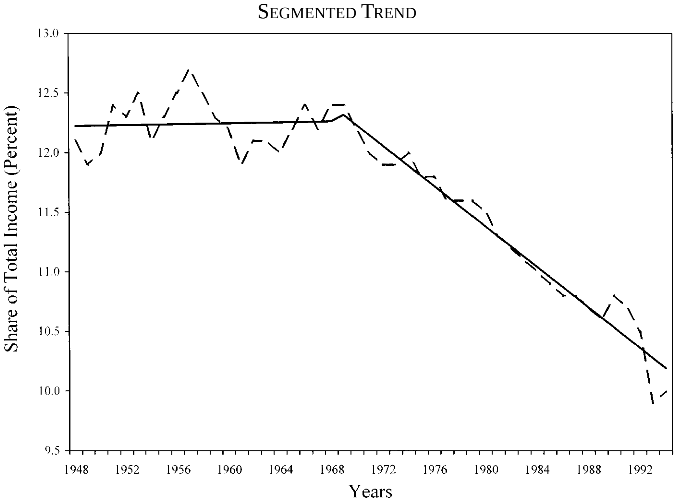


FIGURE 4C.—SHARE OF THE MIDDLE QUINTILE AND ITS ESTIMATED SEGMENTED TREND

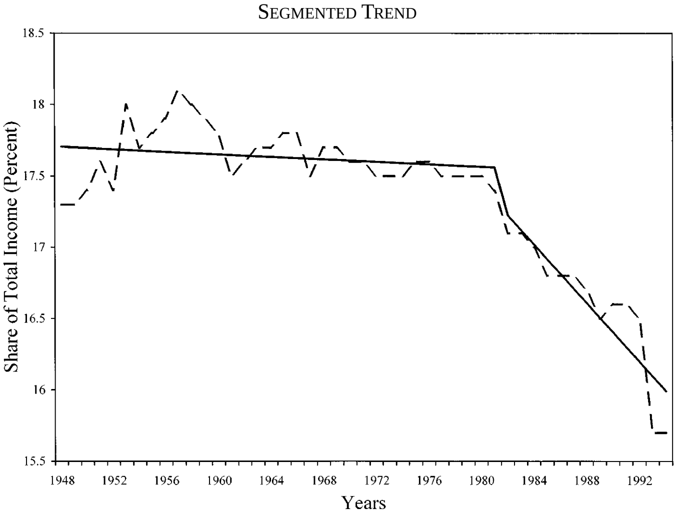


FIGURE 4D.—SHARE OF THE FOURTH QUINTILE AND ITS ESTIMATED SEGMENTED TREND

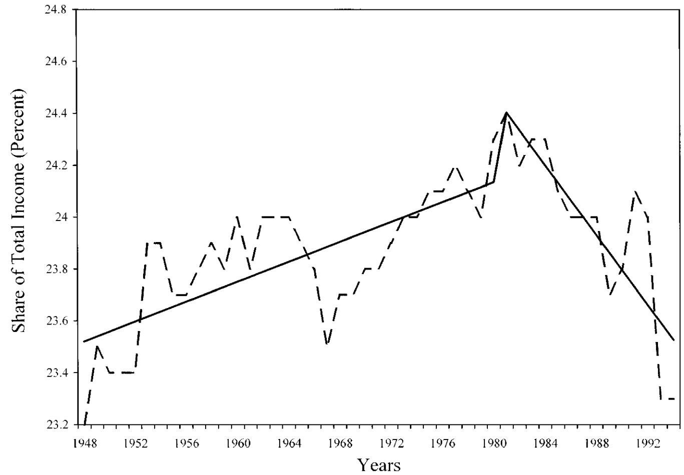
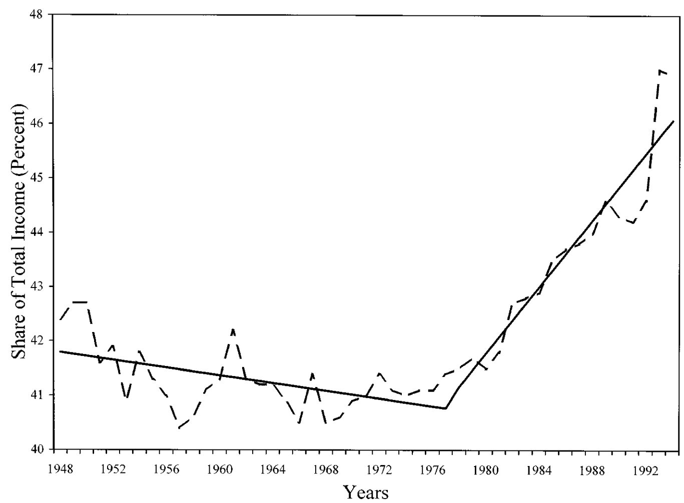


FIGURE 4E.—SHARE OF THE FIFTH QUINTILE AND ITS ESTIMATED SEGMENTED TREND



Hildreth-Lu procedure. Figures 5A, 6A, and 7A display the actual and predicted income shares of the bottom quintile, the middle quintile, and the highest quintile obtained from the traditional and segmented-trend models.[10] The forecasts after 1983 are inaccurate. The forecasts obtained from the traditional model are very similar to the ones reported by Cutler and Katz (1991, p. 15), who show that income inequality is not predicted accurately after 1983, if the coefficients of the explanatory variables are obtained using data that span 1948 to 1983. Figures 5B, 6B, and 7B demonstrate that—when the models are estimated with structural and cyclical unemployment, and using income shares and inflation in differences—very accurate forecasts are obtained. The upward drift in the income share of the highest quintile and the drop in the shares of the lowest and middle quintiles are captured accurately.

[10] The graphs of the other two quintiles were consistent with the ones reported in the paper. To conserve space they are not displayed, but are available upon request.

FIGURE 5A.—ACTUAL AND PREDICTED SHARES MODELS WITH TOTAL UNEMPLOYMENT

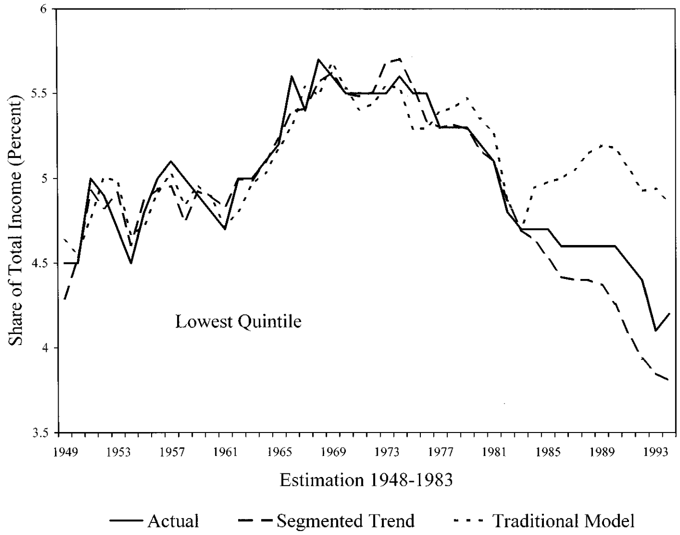
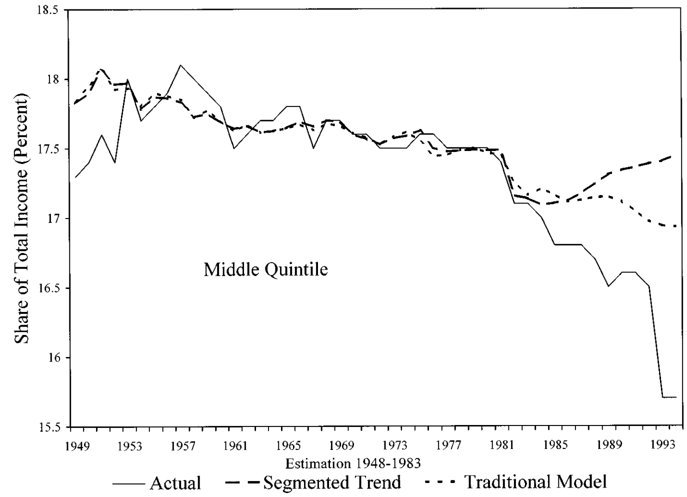


FIGURE 6A.—ACTUAL AND PREDICTED SHARES MODELS WITH TOTAL UNEMPLOYMENT



The contrast between figures 5A–7A and 5B–7B is striking. For example, figure 5A demonstrates that the income share of the lowest quintile is 4.2% in 1994, but the traditional model used by previous research predicts this share as 4.86: an error of 16%. The segmented trend model's forecast is 3.81% in 1994, generating an underprediction of 9%. Figure 5B, on the other hand, shows that the models with structural unemployment predict the share of the lowest quintile accurately. The predictions for 1994 of Models 1 and 2 are 4.06%, and the prediction of Model 3 is 4.08%. The accuracy of the forecasts obtained from the models with structural unemployment in comparison to the ones obtained from total unemployment holds for all income shares. Figure 6A shows that the income share of the middle quintile is 15.7% in 1994. The traditional model with total unemployment generates an overprediction of 1.23% points for the income share of this group in 1994, which amounts to a forecast error of 8%. The segmented trend model predicts the income share of the same group 1.74 percentage points higher than actual. On the other hand, the predictions of Model 1, 2, and 3, presented by figure 6B, are 15.6, 15.7, and 15.7, respectively. Figure 7A displays the actual value of the income share of the highest quintile along with its predicted values obtained from the traditional and segmented-trend models. After 1983, both models underpredict the share of the highest quintile. The actual share is 46.9% in 1994. The forecast of the traditional model is 42.52, and that of the segmented-trend model is 44.63. The forecasts obtained from the three models using structural unemployment, presented in figure 7B, are 47.2, 47.1, and 46.9.

Table 8 presents the mean-squared forecast errors obtained from the models. The first panel displays the mean-squared errors obtained by estimating the models until 1983, and forecasting income shares from 1984 to 1994. The second panel pertains to estimating the models using data

FIGURE 5B.—ACTUAL AND PREDICTED SHARES MODELS WITH STRUCTURAL UNEMPLOYMENT

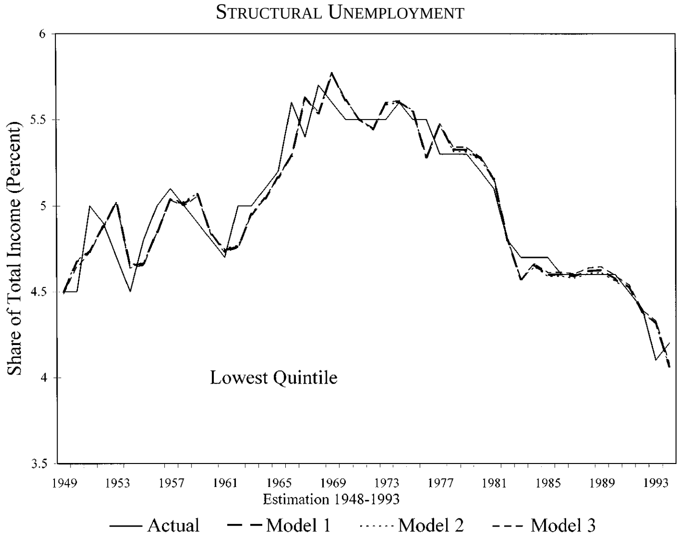


FIGURE 6B.—ACTUAL AND PREDICTED SHARES MODELS WITH STRUCTURAL UNEMPLOYMENT

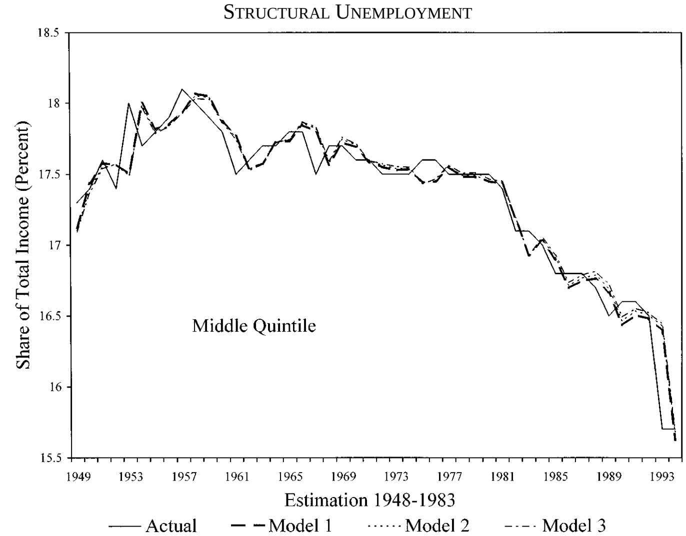


FIGURE 7A.—ACTUAL AND PREDICTED SHARES MODELS WITH TOTAL UNEMPLOYMENT

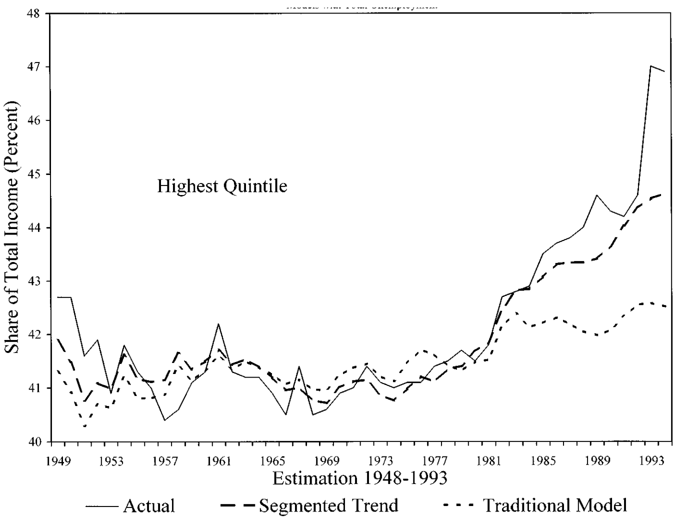


TABLE 8.—PREDICTION PERFORMANCE OF COMPETING MODELS[a]

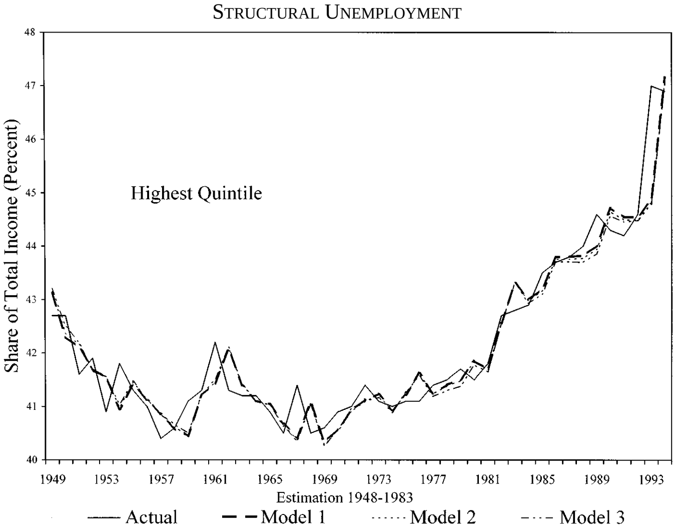
	Traditional Model	Segmented Trends	Model 1	Model 2	Model 3
Mean Squared Forecast Errors for 1984–1994 (Estimation 1948–1983)					
Lowest Quintile	1.475	0.013	0.007	0.007	0.007
Second Quintile	1.357	0.069	0.032	0.032	0.032
Middle Quintile	1.384	0.124	0.053	0.057	0.060
Fourth Quintile	0.403	0.098	0.071	0.074	0.073
Highest Quintile	16.715	1.158	0.487	0.511	0.533
Mean Squared Forecast Errors for 1992–1994 (Estimation 1948–1991)					
	Traditional Model	Segmented Trends	Model 1	Model 2	Model 3
Lowest Quintile	0.188	0.010	0.022	0.022	0.022
Second Quintile	0.364	0.095	0.093	0.095	0.092
Middle Quintile	0.656	0.215	0.167	0.175	0.183
Fourth Quintile	0.448	0.163	0.166	0.172	0.178
Highest Quintile	6.392	1.725	1.521	1.571	1.618

Notes: [a] Model 1 uses StrUr-1 and CcyUr-1 as regressors. Model 2 uses StrUr-HP and CcyUr-HP, and Model 3 uses StrUr-KF and CcyUr-KF.

until 1991, and predicting the income shares for the years 1992 to 1994. The traditional model that regresses the level of income shares on the level of inflation, unemployment, unemployment squared, and a trend produces the highest mean-squared errors. The segmented-trend model performs better than the traditional model, but neither one predicts as good as Models 1 through 3. These results clearly show that the estimated relationship between income inequality and macroeconomic variables for the period 1948–1983 is capable of predicting the income shares of 1984–1994 when the model includes structural and cyclical unemployment as separate regressors, and when it is estimated in differences of income shares and inflation.

In sum, it is evident that families that belong to the bottom sixty percent of the income distribution can improve their relative positions not so much with the help of temporary reductions in unemployment but by a decline in the long-term unemployment rate. Furthermore, a sustained growth in real GNP is not necessarily associated with the well-being of the poor, because a sustained GNP growth may coexist with an increase in the structural unemployment rate, as was the case between 1984 and 1990 in the United States. Rising tides can lift the boat of the poor, with the proviso that long-term unemployment is used as the measure of the tide, rather than the GNP growth.

FIGURE 7B.—ACTUAL AND PREDICTED SHARES MODELS WITH STRUCTURAL UNEMPLOYMENT



V. Summary and Conclusions

This paper investigates the influences of inflation and unemployment on income distribution in the United States. Although it is consistent with the previous work in the choice of the dependent and independent variables, it differs from it in two important ways. First, the paper utilizes recent developments in time-series econometrics that demonstrate that stochastic trend behavior is a better representation of the trends in income shares and inflation than is a deterministic trend. Second, assuming that permanent unemployment might have a different impact on income distribution, the unemployment rate is decomposed into its structural and cyclical components that are used as separate explanatory variables.

Previous research used the unemployment rate as a business cycle proxy and found a significant regressive impact of the unemployment rate on income equality. Decomposition of unemployment into cyclical and structural components reveals that an increase in structural unemployment increases the income share of the highest quintile, and decreases the shares of the bottom sixty percent of the population. In agreement with previous studies that used data from the United States, inflation is found to have a progressive impact on income inequality. Decomposition of inflation into anticipated and unanticipated components revealed that anticipated inflation has no impact on income inequality, but unexpected inflation redistributes income from the highest quintile to the bottom three quintiles.

The results suggest that, although policies that aim to prevent a worsening in income inequality by combatting cyclical downturns have validity, a sustained growth in real output cannot improve income inequality if it is not accompanied by a reduction in long-term unemployment.[11] It is, of course, not obvious how to decrease long-term unemployment. Policies may range from incentives for the employers to hire less-skilled workers, to training programs for the workers who face both stagnant wages and longer spells of unemployment.[12] Despite the debate regarding its effectiveness in the long-run, demand management policies can also be considered to reduce long-run unemployment (e.g., Bean, 1994), especially given the evidence indicating that the unemployment rate depends on its own past, and thus a reduction in current unemployment may help generate a reduction in long-term unemployment (Blanchard & Summers, 1987, 1986).

REFERENCES

Beach, C., and G. Slotsve, ''Recession and Recovery: Men's Earnings Through the 1980s,'' *Journal of Income Distribution* 4:1 (1994), 121–146.

———''Cyclical Sensitivity of Aggregate Income Inequality,'' *Review of Economics and Statistics* 59 (1977), 56–66.

Bean, Charles, ''The Role of Demand Management Policies in Reducing Unemployment,'' in *Reducing Unemployment: Current Issues and Policy Options* (Federal Reserve Bank of Kansas City, 1994) 99–132.

Bishop, John A., J. F. Formby, and R. Sakano, ''Evaluating Changes in the Distribution of Income in the United States,'' *Journal of Income Distribution* 4:1 (1994), 79–105.

Bjorklund, A. ''Unemployment and Income Distribution: Time Series Evidence from Sweden,'' *Scandinavian Journal of Economics* 93 (1991), 457–465.

Black, F., ''General Equilibrium and Business Cycles,'' National Bureau of Economic Research Working Paper, No: 950 (Cambridge, Massachusetts, 1982).

Blackburn, Keith, and Morton O. Ravn, ''Business Cycles in the United Kingdom: Facts and Fictions,'' *Economica* 59 (1992), 383–401.

Blanchard, O. J., and Danny Quah, ''The Dynamic Effects of Aggregate Demand and Supply Disturbances,'' *The American Economic Review* 79:4 (1989), 655–673.

Blanchard, Olivier, and Lawrence Summers, ''Beyond the Natural Rate Hypothesis,'' *American Economic Review Papers and Proceedings* (May 1988), 182–187.

Blanchard, Olivier, and Lawrence Summers, ''Fiscal Increasing Returns, Hysteresis, Real wages and Unemployment,'' *European Economic Review* 31 (1987), 543–566.

Blanchard, Olivier, and Lawrence Summers, ''Hysteresis and the European Unemployment Problem,'' in S. Fisher (ed.), *NBER Macroeconomics Annual* (1986), 15–78.

Blank, R., and A. Blinder, ''Macroeconomics, Income Distribution and Poverty,'' in S. Danziger and D. Weinberg (eds.), *Fighting Poverty* (Cambridge: Harvard University Press, 1986).

Blejer, M. I., and I. Guerrero, ''The Impact of Macroeconomic Policies on Income Distribution: An Empirical Study of the Philippines,'' *Review of Economics and Statistics* (1990), 414–423.

Blinder, A., and H. Y. Esaki, ''Macroeconomic Activity and Income Distribution in the Postwar United States,'' *Review of Economics and Statistics* 60 (1978), 604–609.

[11] A similar point is made by Cutler and Katz (1992).

[12] A detailed discussion of this issue can be found in Katz (1994).

Buse, A., ''The Cyclical Behaviour of the Size Distribution of Income in Canada: 1947–78,'' *Canadian Journal of Economics* 15 (1982), 189–204.

Cutler, David M., and Lawrence F. Katz, ''Macroeconomic Performance and the Disadvantaged,'' *Brookings Papers on Economic Activity* (1991), 1–61.

——— ''Rising Inequality? Changes in the Distribution of Income and Consumption in the 1980s,'' *The American Economic Review,* Papers and Proceedings (May 1992), 546–551.

Danthine, Jean-Pierre, and Michel Girardin, ''Business Cycles in Switzerland: A Comparative Study,'' *European Economic Review* 33 (189), 31–50.

Dickey, D. A., and W. A. Fuller, ''Likelihood Ratio Statistics for Autoregressive Time Series with a Unit Root,'' *Econometrica* 49 (1981), 1057–1072.

Engle, Robert F., and Clive W. J. Granger, ''Introduction,'' in Engle R. F. and Clive W. J. Granger (eds.), *Long-run Economic Relationships* (New York: Oxford University Press, 1991), 1–16.

Flückiger, Yves, and M. Zarin-Nejadan, ''The Effect of Macroeconomic Variables on the Distribution of Income: The Case of Switzerland,'' *Journal of Income Distribution* 4:1 (1994), 25–39.

Fuller, W., *Introduction to Statistical Time Series* (New York: Wiley, 1976).

Granger, Clive W. J., and Paul Newbold, ''Spurious Regressions in Econometrics,'' *Journal of Econometrics* (1974), 111–120.

Harvey, A. C., ''Trends and Cycles in Macroeconomic Time Series,'' *Journal of Business and Economic Statistics* 3 (1985), 216–226.

Harvey, A. C., and P. H. J. Todd, ''Forecasting Economic Time Series with Structural and Box-Jenkins Models,'' *Journal of Business and Economic Statistics* 1 (1983), 299–315.

Hayes, Kathy, D. J. Slottje, S. Porter-Hudak, and G. Scully, ''Is the Size Distribution of Income a Random Walk?'' *Journal of Econometrics* 43 (1990), 213–226.

Hendry, David F., and Adrian J. Neale, ''A Monte Carlo Study of the Effects of Structural Breaks on Tests for Unit Roots,'' in Peter Hackl and Anders H. Westlund (eds.), *Economic Structural Change: Analysis and Forecasting* (Berlin: Springer-Verlag, 1991).

Hodrick, R., and E. C. Prescott, ''Postwar U.S. Business Cycles: An Empirical Investigation,'' mimeo (Carnegie-Mellon University, 1980).

Jäntti, Markus, ''A More Efficient Estimate of the Effects of Macroeconomic Activity on the Distribution of Income,'' *Review of Economics and Statistics* 76 (1994), 372–377.

Jovanovic, Boyan, and Masako Ueda, ''Contracts and Money,'' *Journal of Political Economy* 105 (August 1997), 700–708.

Joyce, Theodore, and H. N. Mocan, ''Unemployment and Infant Health: Time-Series Evidence from the State of Tennessee,'' *The Journal of Human Resources* 28:1 (1993), 185–202.

Katz, Lawrence F., ''Active Labor Market Policies to Expand Employment and Opportunity,'' in *Reducing Unemployment: Current Issues and Policy Options* (Federal Reserve Bank of Kansas City, 1994), 239–290.

Lilien, D. M., ''Sectoral Shifts and Cyclical Unemployment,'' *Journal of Political Economy* 90 (1982), 777–793.

MacKinnon, James, ''Critical Values for Cointegration Tests,'' in R. F. Engle and Clive W. J. Granger (eds.), *Long-run Economic Relationships.* (New York: Oxford University Press, 1991), 1–16.

Macunovich, D. J., and R. A. Easterlin, ''Application of Granger-Sims Causality Tests to Monthly Fertility Data, 1958–1984,'' *Journal of Population Economics* 1:1 (1988), 71–88.

Metcalf, C. E., ''The Size Distribution of Personal Income During the Business Cycle,'' *American Economic Review* (1969), 657–668.

Mocan, Naci, ''Structural Unemployment, Cyclical Unemployment and Income Inequality,'' CRESP Working Paper (University of Colorado at Denver, 1995).

——— ''Is there a Unit-Root in U.S. Real GNP? A Re-Assessment,'' *Economics Letters* 45 (1994), 23–31.

Nelson, C. R., and H. Kang, ''Pitfalls in the Use of Time as an Explanatory Variable in Regression,'' *Journal of Business and Economic Statistics* 2 (1984), 73–82.

Nelson, C. R., and H. Kang, ''Spurious Periodicity in Inappropriately Detrended Time Series,'' *Econometrica* 49 (1981), 741–751.

Nolan, B., ''Macroeconomic Conditions and the Size Distribution of Income: Evidence from the United Kingdom,'' *Journal of Post Keynesian Economics* (1988–1989), 196–221.

Pagan, Adrian, ''Econometric Issues in the Analysis of Regressions with Generated Regressors,'' *International Economic Review* 25(1) (1984), 221–247.

Palmer, J., *Inflation, Unemployment and Poverty* (Massachusetts: D.C. Heath, 1973).

Perron, Pierre, ''The Great Crash, the Oil Price Shock and the Unit Root Hypothesis,'' *Econometrica* 57(6) (1989), 1361–1401.

Phillips, Peter C. B., ''Understanding Spurious Regressions in Econometrics,'' *Journal of Econometrics* 3 (1986), 311–340.

Raj, Baldev, and Daniel Slottje, ''The Trend Behavior of Alternative Income Inequality Measures in the United States From 1947–1990 and the Structural Break,'' *Journal of Business and Economic Statistics* 12(4) (1994), 479–487.

Schultz, T. P. ''Secular Trends and Cyclical Behavior of Income Distribution in the United States: 1944–1965,'' in L. Soltow (ed.), *Six Papers on the Size Distribution of Wealth and Income* (New York: National Bureau of Economic Research, 1969).

Silber, Jacques, and Ben-Zion Zilberfarb, ''The Effect of Anticipated and Unanticipated Inflation on Income Distribution: The Israeli Case,'' *Journal of Income Distribution* 4(1) (1994), 41–49.

Stock, James H., ''Confidence Intervals for the Largest Autoregressive Root in U.S. Macroeconomic Time Series,'' *Journal of Monetary Economics* 28 (1991), 435–459.

Stock, J. H., and Watson, M. W. ''Variable Trends in Economic Time Series,'' *Journal of Economic Perspectives* 2 (1988), 147–174.

Zivot, Eric, and Donald W. K. Andrews, ''Further Evidence on the Great Crash, the Oil-Price Shock, and the Unit-Root Hypothesis,'' *Journal of Business and Economic Statistics* 10(3) (1992), 251–270.