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Is inflation bad for income inequality:

The importance of the initial rate of inflation

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Abstract

This paper explores theoretically and empirically the effects of monetary policy and inflation on income inequality in developed economies. The few empirical studies touching on this issue have given conflicting answers, and the impact on inequality remains puzzling. In an attempt to solve this puzzle we argue that the effects of monetary policy and inflation on inequality depend on the initial rate of inflation. Though in high inflation countries restrictive monetary policy is often beneficial for inequality, reducing inflation in economies with initially low inflation might increase inequality. Empirical investigations for the US and a sample of 15 OECD countries seem to support this hypothesis.

1. Introduction

The purpose of this paper is to investigate the effects of monetary policy and inflation on income distribution. As such it forms part of new research activities on the relation between economic policies, economic growth and poverty, which have been the subject of a flurry of recent publications, mainly in relation to developing countries. Although considerable differences exist among researchers, it is generally accepted that economic policies aiming at stimulating growth need to consider effects on poverty and inequality, by emphasising equitable growth policies and by explicit redistributive policies; the relative importance of these policies depending on the initial economic situation and the socio-political structure in the country in question (Dagdeviren, van der Hoeven and Weeks, 2001).

In order to contribute a different angle to this debate we thought it of interest to investigate the situation in developed countries. In order to keep the research focussed we concentrated on one aspect of economic policy, namely monetary policy. This choice was not only dictated by concerns to arrive at a manageable research agenda, but also by the current debate on the use of monetary policy as an active instrument of economic policy. Researchers and policy-makers both in the United States and in the European Union differ as to the appropriateness of monetary policy in broader policy making. The restrictive stance of the European Central Bank on this issue is well known and criticised by many scholars and international policy-makers, who argue for a broader role of monetary policy in terms of stimulating growth and reducing unemployment. However, given the dis-equalising tendencies in most industrialised countries in the 1990s and the beginning of the current decade, we thought it important not only to look at implication of monetary policies on growth (see for example Brown, 2001) but also to take, as in the case of developing countries, the issue of poverty and inequality into account. We follow Blank and Blinder (1986) in their observations that if one is interested in poverty, a case can be made that the share of income received by the lowest 20 per cent of families or another inequality measure like the Gini ratio is at least as good an index of poverty as the official poverty rate, thus avoiding the complications around the definition of an absolute poverty line. Furthermore, we investigate effects on personal income inequality rather than factor income inequality (as in Argitis and Pitelis, 2001) because the latter, although a major determinant of inequality, is less sensitive for poverty indices.

In looking at the effects of monetary policy and inflation on income distribution this paper takes a different route from the recent stream of literature on the relationship between income inequality and growth. A large debate, in fact, has arisen since the article by Persson and Tabellini (1994) on whether initial inequality is beneficial or detrimental to growth. The debate is still open, with authors piling in both positions. Although related to our topic, this literature differs from the present study in that it looks at the opposite causality from inequality to macroeconomic outcomes, and focuses mainly on fiscal outcome and the consequences for growth; it does not deal with monetary outcomes[1].

The remainder of this paper is organised as follows. Section 2 discusses what we call "the inflation-inequality" puzzle by reviewing previous empirical literature. Section 3 places these findings in a theoretical context examining both the short-run and long-run effects of

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[1] For recent contributions to this debate see Banerjee and Duflo (2000), Bertola (2000), Forbes (2000) and Galor (2000). Two studies tracing explicitly the causality from inequality to monetary policy or inflation are the articles by Beetsma and Van der Ploeg (1996) and by Al-Marhubi (2000).

monetary policy on inequality. Section 4 discusses the outcome of empirical analysis carried out with data we collected from the United States and 15 OECD countries. Section 5 concludes.

2. The inflation-inequality puzzle

The impact of monetary policy and inflation on personal income distribution has been studied in a non-systematic way. A major determinant of the distribution of income in a country is traditionally assumed to be the level of development: as predicted by the so-called Kuznets hypothesis (Kuznets 1955) countries shift from relative equality to inequality and back to greater equality as they move through the development stages. A large number of multi-country empirical studies have shown however that the Kuznets hypothesis explains only a very limited part of the inter-country variation in income distribution (see Bulir and Gulde, 1995, for a review of the literature), and that other policy and structural variables – such as tax and government spending, social transfers, state employment or human capital – improve significantly the explanation of the cross-country differences in income distribution (see among others Milanovic, 1994; Tanzi, 1998; and Chu, Davoodi and Gupta, 2000)[2].

Some of these multi-country studies include inflation among the explanatory variables for income inequality or poverty indicators, but do not have a specific interest in studying the relationship between inflation and income distribution or poverty rates. Few studies are concerned specifically with this question, most of them using time-series data for the United States. The results from all these studies are noticeably mixed – some authors find inflation to be a regressive tax, others find it to be a progressive tax, and others find it to be unrelated to income distribution – so that the literature seems to have generated an inflation-inequality puzzle.

In order to help assessing the inflation-inequality puzzle, tables 1A and 1B summarise the empirical findings of studies known to us on the effects of inflation on different measures of inequality and poverty, divided by nature of the sample (time-series studies in panel A and cross-country/panel data studies in panel B). As shown in table 1A, most time series studies find inflation to be either progressive or non-significant, with the most technically advanced papers in this group (such as Jäntti, 1994 or Mocan, 1999) confirming the progressive effect of inflation on income distribution in the United States over the past five decades. Table 1B shows instead that in all cross country and panel data studies inflation is either regressive or non-significant, but in no case it is progressive.

At first glance the inflation-inequality puzzle stands out clearly. But a more cautious observation suggests a simple way of solving the apparent inflation-inequality puzzle. The effect of inflation on income inequality might depend on the initial level of inflation: when initial inflation is high, reducing inflation might decrease inequality; when initial inflation is low, instead, reducing inflation might come at the cost of higher inequality. This would explain why inflation seems progressive only in time series studies, concerning low-inflation countries, while is found regressive in most cross-country and panel data samples including high-inflation developing economies. There are, however, some exceptions. First, among the cross-section studies, Romer and Romer's (1998) finding of inflation significantly worsening the bottom

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[2] Although originally formulated as a time-series proposition, the Kuznets hypothesis has been assessed mainly in cross-country studies due to the lack of long time-series data on income distribution for a single country.

quintile's average income and inequality in a cross-section of OECD countries in 1988. Given the relatively low level of inflation in the OECD countries in the late 80s, under our hypothesis we would expect inflation to be progressive or at least non-significant. However, Atkinson and Brandolini (1999: p.8) show that Romer and Romer's results for the OECD countries are not robust, and that using strictly comparable data on income inequality (as opposed to the Deininger and Squire (1996) "accept" data set) the inflation variable ceases to be significantly different from zero[3]. In addition, two single-country studies contradict our hypothesis: Power's (1995) finding of United States inflation significantly increasing a consumption-based poverty rate in 1959-89 – but the same author finds inflation not significant with respect to income-based poverty – and Blank and Blinder's (1986) finding of United States inflation increasing poverty in 1959-83 – but the same authors find inflation to increase significantly the lower quintiles income shares. However, more technically advanced studies on United States inequality (such as Mocan, 1999) confirm the progressive impact inflation had in the United States during the last 5 decades.

In this paper we intend to contribute to the solution of the inflation-inequality puzzle by making the explicit assumption that the effect of inflation on income distribution *changes direction* depending on the initial rate of inflation. More specifically we investigate the hypothesis that long-run inflation has a non-monotonic impact on income inequality by estimating a model for inequality with a quadratic term in long-run inflation in two samples (the United States from 1966 to 1999, and a panel of 15 OECD countries from 1973 to 1996). We expect the quadratic term to be positive and significant, suggesting that the relationship between inflation and inequality is U-shaped: income inequality decreases as inflation rises from low to moderate rates, and increases again as inflation grows beyond a certain threshold.

To our knowledge, the hypothesis of a non-monotonic relationship between inflation and inequality has not been tested yet in the literature. In fact, the vast majority of empirical studies assume a linear relationship between income distribution and inflation. Only three studies, among those known to us, assume some form of non-linear (but not necessarily non-monotonic) inflation-inequality relationship.

The first is a study by Bulir and Gulde (1995) testing the impact of inflation on income distribution in time series data for Finland, Israel and Russia. The authors compare their results to previous estimates obtained with similar models for a selection of developed countries. Interestingly, they come to the conclusion that whether inflation is a progressive or regressive tax depends on the level of development and the sophistication of the financial structure: in lower income countries with a relatively unsophisticated financial sector higher inflation tends to be associated to higher inequality, while in richer financially sophisticated countries the poor gain from higher inflation. In other words, Bulir and Gulde conclude that the inflation-inequality relationship is non-monotonic but relate the non-monotonicity to the level of development and the sophistication of the financial markets (not to the level of initial inflation as we do) nor they test the non-monotonic relationship in a cross-country sample.

Second, Easterly and Fischer (2001) estimate the relationship between three measures of improvement in well-being for the poor – the change in the poor's share of national income, the percentage decline in poverty, and the percentage change in real minimum wage – and a non-linear transformation of the inflation rate for a pooled cross-country sample drawn from Deininger and Squire's (1996) data set. They find the well being of the poor negatively

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[3] Note that the comparability problem affecting the OECD data contained in Deininger and Squire's "accept" data set affects all estimates based on this data, such as Dollar and Kraay's (2000).

correlated with the inflation transformation, which implies that higher inflation reduces the well being of the poor with a non-linear (but still monotonic) factor.

Third, Bulir (2001) explores the relationship between income inequality and inflation using a cross-section database containing 75 developing and developed countries. Bulir tests the hypothesis of non-linear effects of inflation on income distribution by representing inflation with a set of dummies, each one representing a different range of inflation (hyperinflation above 300 percent annually, high inflation between 41 percent and 300 percent, low inflation between 5 percent and 40 percent, and very low inflation below 5 percent). Bulir finds that reducing inflation from hyperinflation rates significantly lowers income inequality: the Gini coefficient in high and low inflation countries is lower than in hyperinflation countries by 6 and 9 percentage points respectively. Interestingly, Bulir also finds that further reducing inflation toward very low levels allows a reduction by only 6 percentage points in the Gini coefficient with respect to hyperinflation countries. This implies that moving from low to very low inflation causes inequality to rise by 3 percentage points. Even though Bulir does not focus his attention on the latter result, it is of particular interest to us in that it is consistent with our hypothesis that the inflation-inequality relationship is non-monotonic, and that pushing inflation below a certain threshold makes inequality to worsen again.

In sum, the present study is different from the predecessors in that it tests explicitly the hypothesis of a non-monotonic relationship between inflation and inequality, and it does it for the United States as well as a sample covering 15 developed countries. Most cross-country works on income distribution in fact include both developed and developing countries or developing countries only. By selecting only the developed countries we expect to increase the statistical significance, if it exists, of the U-shaped relationship between inflation and income inequality, whose minimum we expect to be at low levels of inflation, so to shed new light on the impact of monetary policy and inflation on income inequality.

3. Monetary policy and income distribution: the channels of transmission

Textbooks predict that restrictive monetary policy in the short run decreases aggregate income and employment. In an open flexible exchange rate economy, the output and employment loss is reinforced by a real appreciation. The contractionary effect on employment in turn implies a downturn in nominal wage growth, and, through this way, a drop in inflation.

In the long run, the lower-than-expected inflation rate induces workers to adjust downwards their inflation expectations and to bargain lower nominal wage rises, so that a slowdown in labour cost pulls down inflation. Once inflation becomes slower than money supply growth, the increase in real money balances reduces interest rate and leads to an expansion in output and aggregate employment. Accordingly, in an open flexible exchange rate economy, a real depreciation contributes to restore output and employment. The expansive effect ends when expected inflation is fulfilled again, and this happens in correspondence to the initial levels of output, unemployment, real interest rate and real exchange rate. As a result, monetary policy is claimed to be neutral in the sense that it has no long-run effect on level and composition of income, the only long-run effect being on inflation (see for instance Blanchard, 2000).

Textbooks do not analyse the effects of monetary policy on income distribution. However, monetary policy does have an impact on income distribution both in the short run – through the cycle in economic activity generated by the policy change – and in the long run

through the change in inflation reached at the end of the adjustment period. Major causal linkages are depicted in Figure 1 and discussed below.

3.1 The short run

The cyclical downturn produced by a restrictive monetary policy affects income distribution through a number of channels, depicted in the top panel of Figure 1. First, the halt in economic growth caused by the increase in interest rates and the real appreciation adds to unemployment, and this is likely to affect different workers at various degrees. Given that hiring and firing costs are generally higher for skilled than for unskilled workers, unemployment tends to affect mostly low-skilled workers, making income inequality to rise in the short run[4]. The degree of worsening in income distribution depends, among other things, on the sensitivity of investment and consumption to higher interest rate and lower expected demand, and on the elasticity of employment to output fluctuations. In periods of low business confidence and in highly flexible labour markets the impact on inequality is likely to be more serious.

The spurt in unemployment can be reinforced by the presence of downward rigidities in nominal wages, making the slowdown in nominal wage rises to lag behind the decline in inflation brought by the monetary restriction. In this case the growth in real wages reinforces the rise in unemployment (Leidy and Tokarick, 1998). Once again, unemployment is likely to hit mainly low-skilled workers, further deteriorating the impact of monetary restrictions on inequality.

Another way in which monetary policy affects income distribution in the short run is through real interest rates. Since expected inflation is constant by definition in the short run, a decrease in money supply growth leads to an increase in both nominal interest rate and real interest rate. The increase in real interest rate, by making the net borrowers worse off and the net lenders better off, augments income inequality because certainly there are more net lenders at the top than at the bottom of income distribution[5].

Finally, to the limited extent that inflation slows down in the short run following a monetary restriction, lower inflation slows down the loss in purchasing power of non-indexed nominal fixed incomes, such as pensions and transfers. This would reduce income inequality given that the poor receive a larger fraction of their income from transfers than do the remainder of the population.

In sum, as shown in the top panel of Figure 1, restrictive monetary policy has predominantly worsening effects on inequality in the short run. Benefits in terms of income distribution just come from the slowdown in inflation, that helps protecting the purchasing power of non-indexed transfers. However, since the impact of restrictive monetary policy on inflation is limited in the short run, the net effect of restrictive monetary policy in the short run is likely to be a worsening in income inequality.

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[4] Moreover the real appreciation distributes its effects on employment unevenly among sectors, mostly affecting export industries. The impact on income distribution is *a priori* uncertain, depending on the relative concentration of low-skilled labour in export industries. In addition the real appreciation, by reducing the relative price of imported goods, affects the purchasing power of residents with an *a priori* ambiguous impact on income distribution.

[5] If financial intermediaries are more prompt to raise loan rates than deposit rates, the worsening effect on inequality is stronger. See Atkinson (1999, p.18) on the role of financial intermediaries in setting interest rates and affecting household welfare.

3.2 The long run

The main purpose of restrictive monetary policy is to lower inflation in the long run. Lower inflation, in turn, can have an impact on income inequality through various channels, depicted in the bottom panel of Figure 1.

Lower long-run inflation slows down the erosion of money purchasing power, and this can affect income distribution in at least three ways. First, it is usually argued that the poor are less able to protect their living standards from inflationary shocks than the non-poor. Due to the existence of entry barriers in most markets for non-money financial assets, in fact, the poor hold greater proportion of their wealth in cash than do the non-poor disproportionately exposing them to purchasing power erosion by inflation (Ferreira, Prennushi, Ravallion, 1999). Therefore, by slowing down the erosion of monetary financial assets, restrictive monetary policy tends to improve income distribution. Second, lower inflation slows down the erosion of the real value of non-indexed public transfers, such as unemployment benefits and pensions. Since transfer recipients typically belong to the poorest part of the population, this would reduce inequality. Third, since private debt is usually set in nominal terms, a lower inflation rate slows down the erosion of the real value of private debt. If the poor are net nominal debtors, as it is likely, this tend to increase income inequality.

Under specific circumstances, monetary policy can be non-neutral, i.e. can have real effects in the long run. First, restrictive monetary policy can enhance long-run growth by achieving macroeconomic stability and stimulating investment. Although the relationship between growth and income distribution is far from being theoretically and empirically settled (see, for instance, Aghion *et al.*, 1999 and van der Hoeven, 2000), higher long-run growth is typically assumed to improve income distribution (see, for instance, Romer and Romer, 1998 or Dollar and Kraay, 2000[6]).

However, the positive effect of disinflation on growth concerns only countries with initially high inflation, most commonly hyperinflation. In low and moderate inflation economies, inflation is unlikely to create a degree of macroeconomic instability that discourages investment and dampens long-run growth. This idea is largely supported by the empirical literature: Hess and Morris (1996) acknowledge the weak association between growth and low inflation, and Barro (1996) estimates that if inflation falls from 5 percent a year to zero, the growth rate will increase by a mere 0.1-0.15 percent a year (a statistically not significant effect). Fischer (1993) finds that very high inflation is not consistent with sustained growth whereas low inflation is not necessary for high growth, some countries having succeeded in growing over sustained periods with persistent inflation between 15 and 30 percent, typically with the assistance of extensive indexation. Finally Sarel (1996) examines the possibility of non-linear effects of inflation on economic growth, and finds evidence of a significant structural break around 8 percent: inflation below 8 percent has no effect on growth, or may even have a slightly positive effect.

A second possible real impact of non-neutral monetary policy in the long run is on real interest rate. Under neutrality of money, in fact, monetary policy actions have no impact on real interest rate in the long run (the so-called Fisher effect): while real interest rate goes up in

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[6] Dollar and Kraay (2000, page 5-6) assert "reducing government consumption and stabilizing inflation are examples of policies that are 'super-pro-poor'. Not only do both of these raise overall incomes, but they appear to have additional positive effect on the distribution of income, further increasing incomes of the poor. In the case of inflation, this additional distributional effect (…) reflects primarily the reduction of inflation from very high levels."

the short run, in the long run nominal interest rate falls with expected inflation and real interest rate goes back to its original value. Therefore, the short-run negative impact on net borrowers is reversed, making the net borrowers better-off and the net lenders worse-off in the long run. On this ground, monetary policy would have no long-run effects on income distribution through real interest rates. However, when initial inflation is very high, a successful disinflation can lead to a reduction in the risk premium. In this case, real interest rates would be lower in the long run, net borrowers would gain from the disinflation, and net lenders would lose, making income distribution more equitable.

Finally, as noted for the short run, the existence of downward rigidities in nominal wages implies that the slowdown in nominal wage rises lags behind the decline in inflation, leading to higher real wages and higher unemployment. If nominal rigidities are persistent – or even if they are temporary, under certain conditions (see Akerlof, Dickens and Perry, 1996) – there may be a permanent trade-off between inflation and unemployment[7]. Hence, there may also be a permanent trade-off between inflation and income inequality.

The bottom panel of Figure 1 summarises the channels through which restrictive monetary policy affects income inequality in the long run. At first glance, the net impact on inequality seems ambiguous. However, it can be noted that some of these channels (marked with a star in Figure 1) are activated only or mainly in high inflation economies. For obvious reasons, in fact, we can expect the three channels linked to the erosion of money purchasing power (real value of monetary assets, real value of unindexed transfers, and real value of private debt) to have a more important impact on income inequality in high-inflation than in low-inflation economies. Similarly, as already noted, the effects of disinflation on growth and real interest rate occur only when initial inflation is high.

On the contrary, the permanent trade-off between inflation and unemployment is accentuated at very low rates of inflation, since workers are more adverse to actual declines in nominal wages than to reductions in real wages achieved through nominal wage rises smaller than inflation. Akerlof, Dickens and Perry (1996) estimate *via* simulations that while at 2 percent inflation the long-run equilibrium unemployment rate is 6.1 percent, at 1 percent inflation it rises to 6.5 percent, and at zero inflation (price stability) it is as high as 7.6 percent. Because of the existence of downward rigidities in nominal wage, then, reducing inflation from low to lower levels could be more disruptive for unemployment, hence for income inequality, than the traditional theory predicts. This is confirmed also by Wyplosz (2000), who finds evidence for a selected group of five European countries of a non-linear relationship between inflation and steady state unemployment, with unemployment first rising and then decreasing starting from zero inflation.

Taking all that into account, it becomes clear that the net long-run impact of restrictive monetary policy on income inequality is not ambiguous but depends on the initial rate of inflation. In high-inflation countries successful disinflation can bring benefits in terms of income distribution, thanks to a slower erosion in money purchasing power, higher long-run growth and lower real interest rates. On the contrary, in low-inflation countries the distribution benefits from disinflation are likely to be trivial, whereas there is a higher chance of running into a permanent trade-off between inflation and unemployment, which would permanently worsen income inequality.

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[7] The existence of a permanent trade-off between inflation and unemployment is controversial in the literature. While we do not specifically support it, we acknowledge the possibility that this permanent trade-off exists, and consider its impact on income inequality.

4. Data and estimation

In the previous section it was argued that there are theoretical reasons to expect the long-run relationship between inequality and inflation to be non-monotonic. The empirical part of this paper investigates the hypothesised existence of a non-monotonic long-run relationship between inequality and inflation by estimating a quadratic model of the form:

$$G = a + b_1 p + b_2 p^2 + g'X$$

where G is the Gini index, π is the long-run inflation rate and X is a vector of other explanatory variables. We expect $b_1 < 0$ and $b_2 > 0$ meaning that inequality decreases when inflation moves from low to moderate rates, and raises when inflation grows beyond a certain threshold. This hypothesis implies that the inflation elasticity of the Gini coefficient – equal to $b_1 + 2 b_2 p$ – is not constant, but depends on the inflation rate: the elasticity is negative (higher inflation leading to lower inequality) for inflation rates below a threshold $p^* = -b_1/(2 b_2)$, and positive above this threshold. We expect the inflation threshold to be low, and for this reason we restrict our analysis to low-inflation countries – namely the United States and a selection of 15 OECD countries – in order to increase the statistical significance of the non-monotonic relationship, if it exists, between inflation and inequality.

4.1 Time-series analysis of income distribution in the United States

The United States are the only country providing continuous data for income distribution over a long time period, and are therefore the natural object of analysis for time series studies on income distribution. We use data on Gini inequality index and specific quintile income shares from 1966 to 1999 provided by the United States Census Bureau[8], referring to household money income excluding capital gains before taxes.

A visual inspection of the income distribution series (see Figure 2) reveals that inequality in the United States has been rising throughout the period considered, due to a continuous increase in the income share going to the highest quintile at the expenses of the lowest 80 per cent of the distribution[9]. In order to test whether inflation has any long run effect on inequality, running through channels different from growth, we first run a set of six static regressions of the form:

$$g_t = \phi_0 + \phi_1 \pi_t + \phi_2 \pi_t^2 + \phi_3 y_t + u_t \qquad (1)$$

where g is a transformation of the Gini index or of the income share of each quintile[10], π is the long run CPI inflation rate, y is the long run real GDP growth rate and u is an error term[11].

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[8] Data are available on the internet site www.census.gov (Historical Income Tables).

[9] However, in the 60s and 70s the income share of the lowest and the fourth quintiles increased.

[10] The transformation (y* = log (y/(1-y)) is applied to all dependent variables to account for the fact that the Gini coefficient and the income shares lie in the unit interval and to constrain the predicted values to fall between 0 and 1.

[11] An Augmented Dickey Fuller (ADF) test shows that the null hypothesis of a unit root in the Gini index and in the income shares of all quintiles cannot be rejected. However, Raj and Slottje (1994) show that the time series process of income inequality in the United States has experienced a structural break, which biases the ADF test in favour of a unit-root hypothesis. Therefore for several measures of income inequality a segmented trend model is more appropriate than a stochastic trend model. Moreover, given that the values of the Gini …/…

Long-run inflation and growth are obtained through a Hodrick-Prescott filter (with a smoothing parameter 100) from the original series. Estimation is run via ordinary least squares over the sample 1966-1999.

Table 2 presents the estimated coefficients of the static regressions and the implied inflation threshold (the turning point of the U-shaped relationship). In all regressions (except the lowest quintile) the estimated coefficients are highly significant and with the expected signs, and the adjusted R-squared ranges from 0.7 in the second quintile to 0.9 in the fourth quintile. In the Gini regression the coefficient on squared inflation is positive, indicating that starting from zero inflation, the Gini coefficient decreases (higher equality) up to an inflation threshold (estimated around 8 per cent) and increases afterwards. In the quintile income shares regressions the coefficient on squared inflation is negative for the lowest (not significant), second, third and fourth quintiles, and positive for the highest quintile. This indicates that starting from zero inflation, the income share of the lowest 80 per cent of the distribution increases up to an inflation threshold (estimated around 8 per cent) and decreases afterwards. Symmetrically, as inflation rises from zero to about 8 per cent, the income share going to the richest quintile decreases, and rises afterwards. The coefficient on growth is significant in all regressions (except the fourth quintile), positive in the first four quintiles and negative in the Gini and highest quintile regressions, suggesting that growth lowers inequality by increasing the income share of the first 80 per cent of the distribution at the expenses of the richest quintile.

To reduce the finite sample biases in the static regressions we then run an error correction version of the long run relationships given by:

$$g_t = a_0 + a_1 \, g_{t-1} + b_0 \, \pi_t + b_1 \pi_{t-1} + c_0 \, \pi^2_t + c_1 \, \pi^2_{t-1} + d_0 \, y_t + d_1 \, y_{t-1} + \varepsilon_t \quad (2)$$

where the implied long-run coefficients of interest can be recovered as $\phi_1=(b_0+b_1)/(1-a_1)$, $\phi_2=(c_0+c_1)/(1-a_1)$ and $\phi_3=(d_0+d_1)/(1-a_1)$.

Table 3 presents the least squares estimation results for the dynamic equations. The dynamic estimates are generally well behaved: the estimated coefficients on inflation and squared inflation are highly significant in all regressions (except the lowest quintile, and for the first lag of both inflation and squared inflation in the fourth quintile) and the adjusted R-squared scores very high (above 0.9 in all equations). The solved long run coefficients on inflation and squared inflation are all highly significant (except for the lowest quintile) with the expected signs, and imply a threshold inflation rate very close to that estimated with the static regressions, i.e. around 8 per cent.

The long run coefficient on growth is not significant in any of the regressions suggesting that growth had no long run effect on income distribution in the United States in the considered period. Although rather surprising, this is in line with other works finding that growth is not a statistically significant determinant of changes in the distribution of income in cross-country samples (e.g. Easterly and Fischer (2001) and Ravallion and Chen (1997)).

To test for robustness of these results we estimate a number of modifications to the benchmark models presented in equations (1) and (2). First we control for fiscal policy by adding both in the static and the dynamic model (all regressions) the log of the GDP share of government non-defence expenditure. Results for the static and the dynamic regressions are presented in table 2B and 3B respectively. We find that the government expenditure share of GDP is not significant in any of the regressions (except for the lowest quintile static regression

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index and income shares are intrinsically limited in the (0,1) interval, the non-stationarity is conceptually difficult to apply.

where it is only marginally significant), does not alter significantly the results on inflation and growth, nor the estimate of the inflation threshold (which is just slightly lower than in the benchmark equations).

We then run the static and dynamic regressions without the term in squared inflation and present the result in tables 2C and 3C. The linear model in inflation suggests that inflation significantly reduces inequality, by increasing the income share of the lowest 80 per cent of the distribution at the expenses of the richest quintile. Similarly to the quadratic model long run real growth is significant in the static but not in the dynamic regressions, but in both cases long run growth reduces inequality by raising the income shares of the lowest 80 per cent of the distribution, at the expenses of the richest quintile. In all static and dynamic regressions the adjusted R-squared is lower than in the quadratic model (except for the lowest quintile), suggesting that the quadratic model fits the data better.

Next, we test the sensitivity of the coefficients to a change in the sample period used for the estimation. Figure 3 shows the estimated coefficient and the ±2 standard error band from the static quadratic model when the sample is reduced from 1967-99 to 1985-99 by one year at a time. As Figure 3 shows the coefficients of inflation and squared inflation, and the implied inflation threshold, are very stable until 1983, when they suddenly change sign and the quadratic term in inflation is no longer significant. So, after 1983 the quadratic relationship between inflation and inequality is no longer observable, and a linear estimation would yield a negative significant coefficient on inflation indicating that lower inflation increases inequality. This is not at all surprising given that the period 1983-99 contains no observations of inflation above the estimated threshold (8 per cent), so looking at the 1983-99 period allows to observe only the left part of the U-shaped inflation-inequality relationship. Interestingly, however, also the estimated coefficient on growth shows a change after 1983[12], suggesting that year 1983 might mark a structural break in the estimated relationship.

A structural break in the relationship between macroeconomic performance and inequality beginning about 1983 is not new to the literature on inequality in the United States. For instance, Cutler and Katz's (1991) discuss this structural break and the possible alternative explanations for the changing relationship between macroeconomic activity and the income of the poor. They find that increased family income inequality is largely associated with increased wage inequality, particularly of primary earners, and argue that higher wage inequality is related to factors that have shifted relative labour demand away from less skilled workers.

Hence, we estimate the static quadratic model with a dummy for the post-1983 period, and present the results in table 2D. The 1983-99 dummy is highly significant in all but the fourth quintile regression, and shows that inequality was higher in the post-1983 period due to a sharp increase in the top quintile income share at the expenses of the rest of the distribution. The sign and significance of the other coefficients do not change appreciably with the introduction of the time dummy. However, the estimated inflation threshold is remarkably lower than in the benchmark case going down to around 6 per cent. Given that the adjusted R-squared is higher than in the benchmark model for all regressions (except the fourth quintile), the specification presented in table 3D is considered the preferred specification.

In sum the evidence for the United States favour the hypothesis of a non-monotonic long-run relationship between income inequality and inflation, with inequality decreasing as inflation moves away from zero, reaching a minimum in the neighbourhood of 6 per cent

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[12] The coefficient on growth also shows a change – from negative to positive – in 1971 and then stays positive, indicating that higher growth lead to higher inequality after 1971.

inflation, and increasing afterwards. The analysis of the quintile income shares shows that the increase in inequality taking place from zero to 6 per cent inflation is due to the decrease in the income share of the richest quintile, all the other quintiles benefiting from the higher (but still moderate) inflation[13].

4.2 Panel data analysis of income distribution in the OECD countries

Panel data studies of income inequality require a careful selection of the observations to be included in the sample, in order to guarantee the comparability of data across countries and time[14]. We drew our data on income inequality from the World Income Inequality Database (WIID) collected by the World Institute for Development Economics Research (WIDER). The WIID database contains all available information on income inequality over time for developed, developing and transition countries derived from different sources, and classified by income definition (net income, gross income, expenditure or earnings) and reference unit (household, household equivalent, household per capita, family etc.), as well as by area covered (all, rural, urban), population covered (all, employed, taxpayers etc.) and quality ratings (reliable, less reliable).

We first selected all the "reliable" data on the Gini coefficient covering full area (urban and rural) and population for the OECD countries (excluding Mexico, Korea and Hungary who joined the OECD in the 1990s). We then selected the series based on the same reference unit and income definition, in order to have consistent data for cross-country and cross-time comparison. Among all possible combinations (household and gross income, household and net income, household equivalent and gross income etc.), we chose household as the reference unit and net income as the income definition, because this allowed the largest time and country coverage. Since this sub-sample still included overlapping observations from different sources for a single year in some countries, we further selected the data by keeping only the observations from the same source within each country (but different sources across countries). In choosing our preferred source, we looked for the most updated and longest time series[15]. Finally we excluded countries with only one observation (Austria, Luxembourg, Spain, Switzerland and Turkey). This selection process left us with a panel of 88 observations covering 15 OECD countries over the period 1960-1996, shown in Table 4. The number of observations over time for each country varies from 2 (Denmark, Finland and Portugal) to 22 (Italy). In order to make the panel less unbalanced we truncated the sample in 1973 (the first year in which we have observations on at least four countries) and adjusted the sample by selecting only observations two years apart from each other[16]. We then estimate the benchmark regressions over the period 1973-96 for the adjusted sample (60 observations), and test for sensitivity of results to sample variations.

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[13] In the case of the lowest quintile, our estimates indicate a monotonic and positive long run relationship with inflation: the income share of the lowest quintile rises with inflation at all levels of inflation.

[14] Atkinson and Brandolini (1999) warn about combining in cross-country and time series analysis income inequality data based on different income concepts, sampling, demography etc.. They show that the widely used "accept" sub-sample of Deininger and Squire's (1996) data set contains inequality data based on different income concepts and reference units, and can therefore lead to misleading conclusions even if dummy variables are introduced to account for the difference in definitions.

[15] More precisely for each country we selected, first, the source whose last observation is the most recent, and, when two different sources extended to the same year in time, we selected the longer one. We also excluded all sources providing only one observation over time. Choosing the most updated time series has the side-advantage of maximizing the number of observations on income inequality in years of low inflation.

[16] For each couple of adjacent observations, we select the most recent one and exclude the other.

Our indicator for inflation is the yearly percentage variation of the consumer price index (all items) contained in the OECD Main Economic Indicators database. The inflation data (see Table 5) show that average inflation in 1973-1996 for the selected OECD countries is 7.4 per cent, ranging from 3.7 per cent in Western Germany to 16 per cent in Portugal. The median observation on inflation is 6.3 per cent, maximum inflation is found in Portugal in 1977 (33 per cent) and minimum deflation in the Netherlands in 1987 (-0.7 per cent). By containing observations on moderate and low inflation only, our sample should allow to pick up any possible upsurge of inequality at low levels of inflation. Growth data are based on real GDP (Laspeyres index, 1985 international prices) provided by the Penn World Tables 5. Table 6 shows the descriptive statistics of growth, reporting an average annual growth rate of 2.6 per cent, ranging from 4.4 per cent in Ireland to 1.54 per cent in Sweden.

Since our objective is to study the relationship between inflation and income inequality in the log-run, we derive a time series for long-run inflation and long-run real GDP by applying the Hodrick-Prescott filter, with a smoothing parameter 100, to the original data on inflation and real GDP of each country. We then estimate via the fixed effect estimator the following regression:

$$G_{it} = a_i + b_1 p_{it} + b_2 p_{it}^2 + g y_{it} + e_{it}$$

where G_{it} is the Gini index in country i at time t, a_i is a country-specific fixed effect, π_{it} is long-run inflation, y_{it} is long-run real GDP growth, and e_{it} is a standard normal error[17]. Note that we choose not to include per capita income and its square among the regressors because the Kuznets hypothesis is supposed to explain the variability of inequality along the development process, whereas the OECD countries in the period considered were largely in the same stage of development. Moreover, the Kuznets hypothesis is not widely supported by the empirical evidence. In addition, note that by including growth among the regressors, the effect of inflation on inequality working through overall growth is captured there. The coefficients on inflation and squared inflation then capture any additional redistributive effect inflation can have other than through growth.

The first column of table 7 shows the estimated coefficients and the White heteroskedasticity-consistent standard errors of the benchmark model. The coefficients on long-run inflation rate and squared rate have the expected sign – negative and positive respectively – and are highly significant. The estimated coefficient on long-run growth is highly significant and positive, which is somewhat surprising indicating that higher growth leads to higher inequality in the considered sample. The fit of the regression is relatively good (the adjusted R-squared is 0.82).

Using the long-run inflation and squared inflation estimated coefficients, $\widehat{b}_1$ and $\widehat{b}_2$, the value of the inflation threshold minimising income inequality can be calculated as $\widehat{p}^* = -\widehat{b}_1 / (2\widehat{b}_2)$. The estimates of $\widehat{b}_1$ and $\widehat{b}_2$ reported in the first column of table 7 imply an inflation threshold just above 12%. In other words, as inflation grows above zero, income inequality decreases, reaches a minimum around 12 per cent inflation, and rises again for higher

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[17] The fixed effects estimator for panel data allows for country-specific intercepts, while assuming common slope coefficients across countries. The model was estimated also for a transformation of the Gini index, namely log (Gini/(1-Gini)), to account for the fact that the Gini coefficient is limited to the (0, 1) interval. Since the results obtained do not differ qualitatively, we present the results obtained with the Gini index as dependent variable, because the estimated coefficients are more readily interpretable as elasticities.

inflation rates. The change in income inequality caused by a one percentage point change in inflation is represented by the elasticity of the Gini index with respect to inflation, and could be called the Gini sacrifice ratio analogously to the "sacrifice ratio" referring to the unemployment cost of disinflation. This elasticity, represented by $\widehat{b}_1 + 2\widehat{b}_2 p_{it}$ in our model, varies with the initial rate of inflation and is therefore country and year specific.

Table 8 presents the estimated inflation elasticities of the Gini index for alternative values of initial inflation, ranging from 1 to 20%. As shown, lowering inflation by one percentage point when initial inflation is, say, 3%, would increase income inequality by nearly 0.7 Gini points. The Gini sacrifice ratio would be halved (0.33 Gini points) at 8% initial inflation, and would be null at 12% inflation. The elasticity becomes larger and positive for initial inflation rates above 12%, indicating, for instance, a 0.5 Gini points gain in income equality when inflation is reduced from 20 to 19%.

The estimated values of the Gini sacrifice ratio may appear low, but are quite substantial compared to the average value of the Gini coefficient in the OECD countries. Consider for instance the hypothetical income distribution represented in table 9 (case 1). This distribution is not far from the United Kingdom's or United States' distribution of income in the last decade. The corresponding Gini coefficient is 33.2. Transferring income directly from the top to the bottom quintile, so to increase the bottom quintile income share by 10%, would decrease the Gini coefficient by 0.8 (see case 2). This is to show that the apparently low values of the inflation elasticities presented in table 8 may be associated to important changes in the distribution of income.

The stability of estimates of the benchmark model presented in the previous section was tested in a number of different ways.

The second column of table 7 shows the estimates of the benchmark equation without squared inflation. In the linear model the coefficient on inflation is negative (suggesting that higher inflation goes with lower income inequality) but not significant, the coefficient on growth remains positive and significant, and the adjusted R-squared (0.79) is lower than in the quadratic model indicating that the quadratic model fits the data better.

The third column of table 7 reports the estimated coefficients when the GDP share of government expenditure (lagged one year) is added to the benchmark model. The coefficient on fiscal policy is negative – as expected – and significant, leaving the non-monotonic inflation-inequality relationship significant and substantially unchanged (the inflation threshold would be slightly lower than in the benchmark model).

The fourth column of table 7 shows the sensitivity of the benchmark equation to the inclusion of a dummy variable for the high-inflation period 1973-81. The time dummy does not affect the sign or the significance of any of the coefficients of interest, and is positive and significant indicating that in the high-inflation years there was an upsurge of inequality, as predicted by the theory. However, we prefer not to include this dummy in the benchmark model since it picks up the effect of high inflation on inequality, also controlled by squared inflation.

Next, the benchmark model is tested for robustness to variations in the sample. First, each country in turn is excluded from the sample, and table 10 presents the estimated coefficients and inflation thresholds obtained in each case. As shown in the table, the coefficients on inflation and squared inflation are remarkably stable, as it is consequently the implied inflation threshold. Also the coefficient on growth is quite stable, except for the exclusion of Italy and Norway, in which cases the estimate is lower..

   The benchmark model is then tested for robustness to variations in the sample period. Figure 4 shows that restricting the sample period progressively from 1960-96 to 1980-96 has no relevant effect on the value or the significance of the estimated coefficients on inflation and squared inflation, with the inflation threshold stable around 12 per cent. Further reducing the sample to 1981-96 makes the coefficients on inflation and squared inflation jointly non-significant. Similarly the inflation threshold is stable and significant when years prior to 1980 are included, and becomes non-significant if only post-1980 years are included. This is not surprising because since 1984 long-run inflation has been below the estimated 12 per cent threshold in all considered countries (except Portugal). Naturally, when initial inflation for all observations is below the threshold only the decreasing side of the U-shaped inflation-inequality relationship can be observed. In fact, fitting a linear model in long-run inflation and growth over the period 1981-96 gives a negative and significant coefficient on inflation, suggesting that lower inflation increases inequality. Finally note that the coefficient on growth is always positive, but non-significant when years 77-80 are excluded from the sample. From 1981 onwards it is much lower, positive, and (marginally) significant.

   In sum, the evidence presented for the OECD countries supports the hypothesis of a U-shaped long-run relationship between income inequality and inflation. Controlling for long-run growth and for country-specific fixed effects it is found that the lowest inequality is reached around 12 per cent inflation. Pushing inflation below 12 per cent, as well as allowing inflation to go above 12 per cent, makes inequality progressively higher. The 12 per cent inflation threshold found in the OECD countries is higher than that found for the United States, showing a minimum in the Gini coefficient at around 6 per cent inflation. The difference may be partially explained by the lower precision of the estimates obtained from the multi-sources discontinuous OECD panel with respect to the United States time series. However, the result suggests that reducing inflation from moderate (around 12 per cent) to low (around 6 per cent) rates would increase inequality in the OECD (mainly European) countries, whereas it would decrease inequality in the United States.

   Recalling that one important theoretical reason for expecting lower inflation to increase inequality in the long run is the presence of downward stickiness of nominal wages (paragraph 3.2), one could argue that the different result for the 15 OECD countries and the United States may be explained by the different degree of downwardly rigidity of nominal wages in Europe and in the United States. Given that nominal wages are downwardly more rigid in Europe than in the United States, when inflation is reduced from moderate to low levels, nominal wage cuts lag behind the decline in inflation and lead to higher real wages, unemployment and inequality more in Europe than in the United States.

   We would like to stress that these results can only be indicative, given that the scarcity of comparable data on inequality makes it difficult to test any hypothesis on income inequality in cross-country studies to a high degree of confidence. We are also aware that we are explaining only a minor part of the cross-country variability in income inequality, the major part being explained by the country-specific fixed effects. However, the objective of this paper is not to give an exhaustive explanation of the cross-country variability in income inequality, but rather to shed light on the effects of monetary policy and inflation on income distribution in low inflation countries. Moreover, given that considerable attention was put to include in the OECD panel only directly comparable data, and given that the hypothesis of a non-monotonic inflation-inequality relationship is not rejected in two different samples (the United States and the 15 OECD) overall we think that the U-shaped inflation-inequality relationship emerges as a relatively robust finding.

5. Conclusions

This paper is concerned with the effects of monetary policy and inflation on income inequality. As explained in section 2, in the literature the empirical evidence on the relationship between inflation and inequality is mixed, with examples of both positive and negative correlation, producing what we have called the inflation-inequality puzzle. We suggest to solve this puzzle by assuming a non-monotonic relationship between inflation and inequality. Particularly, we suggest the use of a U-shaped relationship between income inequality and inflation, with inequality decreasing as inflation moves from high to low rates, and increasing as inflation is further reduced from low to lower rates.

The theoretical discussion of section 3 has identified a set of possible channels through which monetary policy can affect income inequality in the short and in the long run. The theory suggests that while in the short run restrictive monetary policy can be expected to deteriorate income distribution unambiguously, in the long run the net impact can be different depending on the initial rate of inflation. Particularly, when initial inflation is high the beneficial effects of disinflation on income distribution can be expected to dominate, whereas when initial inflation is low the detrimental effects of disinflation on income distribution may prevail. Therefore the economic theory suggests that the long-run relationship between inequality and inflation is U-shaped.

In section 4 we explored the empirical evidence for the non-monotonic long-run relationship between inequality and inflation in the United States over the period 1967-1999 and in 15 OECD countries using a panel of 60 observations over the period 1973-1996. From both samples we found evidence in favour of a U-shaped long-run relationship between inflation and inequality, with the estimated minimum in inequality around 6 per cent inflation in the United States, and around 12 per cent in the OECD countries. In the case of the United States, where data are available also on the quintiles income shares, the increase in inequality below 6 per cent inflation arises from the richest quintile gaining income share at the expenses of the rest of the distribution.

Unfortunately, the scarcity of cross-country and cross-time comparable data on inequality makes it difficult to estimate any hypothesis about the factors driving income distribution to a high degree of confidence. Hopefully in the future more continuous data will be collected on income distribution for the OECD countries and especially the European Union, so that it will be possible to analyse the topic with higher precision and confidence. We intend to do that in the future, as well as to extend the analysis to relative poverty data. However, even in the absence of extensive data sets, policy makers in the developed countries should take into account the negative permanent effects that restrictive monetary policy could have on income distribution at low inflation rates, and try to minimise them through effective labour and fiscal policies. As a recent contribution of the World Bank (Ferreira *et al.*, 1999) puts it: "macroeconomic policies should aim to achieve stabilisation goals at the least cost to the poor, … and as soon as a sustainable external balance has been reached and inflationary pressures have been contained, macroeconomic policy should be eased (interest rates reduced and efficient public spending restored), to help offset the worst effects of the recession on the poor". These words are referred to the developing countries: our wish is that they could be heard in the developed countries too.

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Table 1 A - The inflation-inequality puzzle (single country studies)

Authors	Sample	Dependent variable	Control variables	Inflation		
				Regressive	**Progressive**	**Not sign.**
Mocan (1999)	US 1948-94	Change in quintile shares of family income	Change in inflation, levels of structural and cyclical unemployment		X	
Johnson and Shipp (1999)	US 1980-94 (quarterly data)	Gini indexes based on pre-tax cash income and consumption-expenditures	Inflation, unemployment, per capita real transfers, lagged Gini, time trend		X	
Romer and Romer (1998)	US 1969-94	Change in poverty rate	Inflation		X	
			Inflation, unemployment			X
		Change in Gini index	Inflation, unemployment			X
		Change in income share of bottom quintile	Inflation, unemployment		X	
Cole and Towe (1996)	US 1947-93	Gini index	Inflation, lagged Gini, GDP per capita, unemployment, time trend			X
Powers (1995)	US 1959-92	Change in income poverty rate	Change in inflation and unemployment			X
		Change in consumption poverty rate	Change in inflation and unemployment	X		
Jäntti (1994)	US 1948-89	Quintile shares of family income	Inflation (GNP deflator), prime-age male unemployment rate and squared rate, time trend, dummy 1981.		X	
Cutler and Katz (1991)	US 1959-89	Poverty rate	Inflation, poverty line/mean or median income, unemployment, post-1983 trend		X	
			Also controlling for lagged poverty rate			X
	US 1947-89	Quintile income shares	Inflation, real per capita GNP, unemployment, post-1983 trend, lagged dep. var.			X
Blank and Blinder (1986)	US 1948-83	Quintile shares of family income	Inflation, unemployment (male 25-54), time trend, lagged dep. var.		X (significant for second quintile)	
	US 1959-83	Poverty rate	Inflation, unemployment, transfers/GDP, poverty line/mean income, lagged poverty rate	X		
Blinder and Esaki (1978)	US 1947-74	Quintile shares of family income	Inflation, unemployment, time trend		X	
Other countries						
Bulii and Gulde (1995)	Finland 1977-84	Gini index	Inflation, unemployment			X
	Israel 1986-92	Gini index	Inflation, unemployment	X		
	Russia 1991-94 (quarterly data)	Gini index	Inflation, unemployment			X
Brandolini and Sestito (1994)	Italy 1977-91	Gini index	Inflation, unemployment, trend		X	
Yoshino (1993)	Japan 1963-88	Decile income shares of worker's households	Expected and unexpected inflation, ratio of job offers to applicants, terms of trade, lagged dependent variable	X(*)		

Regressive (progressive) means inflation increases (decreases) inequality or poverty, or decreases (increases) the lower quintiles income shares.

(*) Both expected and unexpected inflation are estimated to be regressive tax. Expected inflation is insignificant for the lowest 20 and 10 percent groups. Unexpected inflation is insignificant for the 80-51 percent group.

Table 1 B - The inflation-inequality puzzle (cross-country and panel data studies)

Authors	Sample	Dependent variable	Control variables	Inflation		
				Regressive	Progressive	Not sign.
Bulir (2001)	Cross section: various years, 75 dev.ed and dev.ing countries	Gini index (Milanovic data)	4 inflation dummies, per capita GDP, squared per capita GDP, % state employment, transfers/GDP, M2/GDP	non-monotonic (U-shaped)		
Easterly and Fischer (2001)	panel: 110 obs; not specified number of countries; 3 decades	Change in income share of the bottom quintile (DS data set)	Inflation, growth	X		
	panel: 64 obs; 42 dev.ing and transition countries; 1981-93.	Change in poverty rate (50 percent of initial mean income) (Ravalllion and Chen, 1997)	Inflation, growth	X		
Dollar and Kraay (2000)	panel: 232 obs; 80 dev.ed and dev.ing countries; 4 decades	Average income of the bottom quintile (augmented DS high quality data set)	Inflation, per capita income			X
			Inflation, per capita income, openness, government consumption, property rights	X		
Chu, Davoodi and Gupta (2000)	panel: 10 dev.ing and transition countries in 80s and 90s	Gini index (DS data set)	Inflation, ratio of direct to indirect tax, secondary school enrolment and urbanization rates			X
Romer and Romer (1998)	Cross section: 1988, 66 dev.ed + dev.ing countries	Average income of bottom quintile (DS high quality)	Inflation	X		
			Inflation, demand variability			X
	Cross section: 1988, 76 dev.ed + dev.ing countries	Gini index (DS high quality)	Inflation, continent dummies and demand variability	X		
			Inflation, demand variability			X
	Cross section: 1988, 19 OECD countries	Average income of bottom quintile (DS high quality)	Inflation, demand variability	X(*)		
		Gini index (DS high quality)	Inflation, demand variability	X(*)		
Sarel (1997)	Cross-section: various years, 45 dev.ed and dev.ing countries	Change in Gini index	Inflation, plus a large number of macroeconomic and demographic variables			X
Bulir and Gulde (1995)	Panel: 121 obs; 18 dev.ed and dev.ing countries; 1960-92	Gini index	Inflation, per capita GDP, squared per capita GDP, public expenditure/GDP, country-specific dummies	X		
	Panel: 55 obs; 18 dev.ed and dev.ing countries; 1960-92	Change in Gini index	Change in inflation, per capita GDP and public expenditure; country-specific dummies			X

Regressive (progressive) means inflation increases (decreases) inequality or poverty, or decreases (increases) the lower quintiles income shares.

DS data set = Deininger and Squire data set.

(*) Atkinson and Brandolini (1999: p.8) show that using strictly comparable data for the same countries, inflation would be non-significant

Table 2 - Static regressions for the US(*)						
	GINI	**lowest**	**second**	**third**	**fourth**	**highest**
Constant	**0.848354**	**-3.77974**	**-3.30453**	**-2.50273**	**-1.295984**	**0.905282**
	(0.2242)	(0.1245)	(0.2076)	(0.1492)	(0.0470)	(0.1881)
Inflation	**-12.0302**	**4.216041**	**8.61777**	**7.981081**	**4.833194**	**-11.5417**
	(2.4619)	(1.3670)	(2.2799)	(1.6383)	(0.5160)	(2.0657)
Squared Inflation	**72.85701**	-6.32323	**-49.87765**	**-47.9431**	**-33.13026**	**68.41693**
	(23.9950)	(13.3238)	(22.2218)	(15.9682)	(5.0291)	(20.1335)
Growth	**-25.9043**	**13.32523**	**26.74355**	**20.09005**	0.357091	**-23.3224**
	(7.4625)	(4.1437)	(6.9110)	(4.9661)	(1.5641)	(6.2615)
Inflation threshold	**8.26%**	33.34%	**8.64%**	**8.32%**	**7.29%**	**8.43%**
F-statistic	(57.0092)	(0.3118)	(27.3703)	(53.9977)	(393.5048)	(67.3589)
Probability	(0.0000)	(0.5808)	(0.0000)	(0.0000)	(0.0000)	(0.0000)
Adjusted R-squared	0.796017	0.884627	0.721598	0.797784	0.917947	0.845922
F-statistic	42.62537	82.78669	28.64730	43.08224	120.3307	59.56241
Prob(F-statistic)	(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0000)

(*) Dependent variable is log(y/(1-y)) where y is the Gini coefficient, or the lowest, second,
third, fourth and highest quintile income share respectively.
Bold = significant at 5%. Standard errors in brackets.

Table 3 - Dynamic regressions for the US(*)						
	GINI	lowest	second	third	fourth	highest
Constant	0.1182	**-2.4319**	**-1.6006**	**-1.1772**	**-0.8382**	0.2004
	(0.1200)	(0.6918)	(0.4050)	(0.3385)	(0.2383)	(0.1210)
Lagged dependent variable	**0.3922**	0.3099	**0.3816**	**0.4188**	*0.3365*	**0.4315**
	(0.1751)	(0.1925)	(0.1611)	(0.1755)	(0.1884)	(0.1733)
Inflation	**-23.2158**	**9.7028**	**22.5294**	**15.3572**	**4.7679**	**-20.0888**
	(7.6029)	(5.3498)	(6.6548)	(5.1494)	(2.0837)	(6.8238)
Squared inflation	**151.9829**	-39.8000	**-149.0894**	**-100.2629**	**-39.7269**	**133.5582**
	(52.9232)	(41.3542)	(45.3929)	(35.3913)	(17.3445)	(49.0634)
Inflation(-1)	**17.6267**	-7.2798	**-19.0302**	**-11.8143**	-1.8646	**15.0989**
	(6.3035)	(5.5013)	(5.7573)	(4.2031)	(2.0433)	(5.6874)
Squared inflation(-1)	**-117.4339**	36.8537	**128.2437**	**78.3472**	20.7945	**-104.5823**
	(44.9116)	(43.4141)	(39.6074)	(29.3666)	(16.2755)	(42.2578)
Real GDP growth	**16.7772**	0.6698	*-11.2619*	-7.6069	*-6.8604*	12.5898
	(7.9055)	(9.1401)	(5.4853)	(4.5428)	(3.5896)	(7.5651)
Real GDP growth(-1)	**-21.2537**	3.3788	**15.2236**	**11.3207**	*6.5459*	**-17.0550**
	(7.5396)	(8.7260)	(5.2923)	(4.4390)	(3.3763)	(7.3084)
Adjusted R-squared	0.9690	0.9212	0.9742	0.9765	0.9427	0.9686
F-statistic	139.2771	52.7489	168.3626	185.2753	73.8960	137.8210
Prob(F-statistic)	0.0000	0.0000	0.0000	0.0000	0.0000	0.0000
Implied Long Run Coefficients						
Inflation	**-9.1957**	*3.5110*	**5.6583**	**6.0960**	**4.3757**	**-8.7768**
F-statistic	(24.4916)	(3.5427)	(18.6226)	(29.9495)	(33.5275)	(20.4632)
Probability	(0.0000)	(0.0720)	(0.0002)	(0.0000)	(0.0000)	(0.0001)
Squared inflation	**56.8430**	-4.2692	**-33.7083**	**-37.7092**	**-28.5347**	**50.9668**
F-statistic	(10.2594)	(0.0591)	(7.2719)	(12.4810)	(15.6663)	(7.5649)
Probability	(0.0038)	(0.8099)	(0.0126)	(0.0017)	(0.0006)	(0.0111)
Growth	-7.3652	5.8665	6.4062	6.3901	-0.4740	-7.8540
F-statistic	(1.3315)	(0.0753)	(1.9380)	(2.2850)	(0.0307)	(1.3652)
Probability	(0.2599)	(0.7861)	(0.1767)	(0.1437)	(0.8624)	(0.2541)
Inflation threshold	**8.09%**	41.12%	**8.39%**	**8.08%**	**7.67%**	**8.61%**
F-statistic	(69.9032)	(0.0753)	(41.7680)	(78.3583)	(118.2540)	(43.5665)
Probability	(0.0000)	(0.7862)	(0.0000)	(0.0000)	(0.0000)	(0.0000)

(*) Dependent variable is log(y/(1-y)) where y is the Gini coefficient, or the lowest, second, third, fourth and highest quintile income share respectively.

Bold = significant at 5%. Italic = significant at 10%. Standard errors in brackets.

Table 2B - Static regressions for the US controlling for fiscal policy (*)						
Estimation results	**GINI**	**lowest**	**second**	**third**	**fourth**	**highest**
Constant	0.283176	**-3.30057**	**-2.76164**	**-2.09633**	**-1.402949**	0.468684
	(0.5121)	(0.2740)	(0.4733)	(0.3392)	(0.1079)	(0.4314)
Inflation	**-12.4572**	**4.578069**	**9.02794**	**8.288133**	**4.752379**	**-11.8716**
	(2.4656)	(1.3193)	(2.2788)	(1.6334)	(0.5193)	(2.0773)
Squared Inflation	**78.99952**	-11.531	**-55.778**	**-52.36**	**-31.96774**	**73.16199**
	(24.3130)	(13.0097)	(22.4704)	(16.1070)	(5.1209)	(20.4835)
Growth	**-29.6558**	**16.50586**	**30.34716**	**22.78768**	-0.35292	**-26.2204**
	(8.0075)	(4.2847)	(7.4006)	(5.3048)	(1.6866)	(6.7462)
Gov.Expenditure (% GDP)	-0.34536	*0.292802*	0.33174	0.248338	-0.065362	-0.26679
	(0.2819)	(0.1508)	(0.2605)	(0.1868)	(0.0594)	(0.2375)
Inflation threshold	**7.88%**	19.85%	**8.09%**	**7.91%**	**7.43%**	**8.11%**
F-statistic	(71.0534)	(1.3916)	(38.1288)	(69.6478)	(311.8994)	(79.1373)
Probability	(0.0000)	(0.2481)	(0.0000)	(0.0000)	(0.0000)	(0.0000)
Adjusted R-squared	0.799481	0.894679	0.72744	0.803003	0.918542	0.847301

(*) Dependent variable is $\log(y/(1-y))$ where y is the Gini coefficient, or the lowest, second, third, fourth and highest quintile income share respectively.

Bold = significant at 5%. Italic = significant at 10%. Standard errors in brackets.

Table 3B - Dynamic regressions for the US controlling for fiscal policy (*)						
Estimation results	**GINI**	**lowest**	**second**	**third**	**fourth**	**highest**
Constant	-0.1264	**-2.4778**	**-1.5264**	**-1.0028**	**-0.6396**	-0.1577
	(0.4494)	(0.7511)	(0.4374)	(0.3551)	(0.2673)	(0.4256)
Lagged dependent variable	*0.3532*	0.2814	*0.3285*	*0.3342*	0.2560	*0.3659*
	(0.1906)	(0.2056)	(0.1736)	(0.1873)	(0.1944)	(0.1895)
Inflation	**-25.6499**	9.8668	**25.2554**	**19.0181**	**6.5434**	**-23.6941**
	(8.8812)	(6.1865)	(7.6393)	(5.8955)	(2.3417)	(8.1027)
Squared inflation	**178.5528**	-44.0275	**-176.2174**	**-138.7909**	**-65.1659**	**173.9361**
	(71.0374)	(58.0596)	(58.6969)	(45.8386)	(23.2618)	(67.7149)
Inflation(-1)	**19.6383**	-7.6962	**-21.4763**	**-14.8238**	-3.0998	**18.0390**
	(7.3063)	(6.1084)	(6.5767)	(4.7707)	(2.1557)	(6.6247)
Squared inflation(-1)	**-140.3655**	44.6571	**153.2254**	**111.1641**	**40.9544**	**-139.5450**
	(59.7224)	(56.8052)	(50.8518)	(37.7619)	(20.3179)	(56.9795)
Real GDP growth	*18.7032*	3.7996	*-11.8260*	*-10.7570*	**-10.5072**	15.1525
	(9.9603)	(10.9037)	(6.7298)	(5.6307)	(4.3638)	(9.4563)
Real GDP growth(-1)	**-25.3134**	2.0530	**18.0205**	**17.4001**	**12.2216**	**-23.1119**
	(11.4981)	(12.7997)	(7.9434)	(6.6924)	(4.8925)	(11.1898)
Gov.Expenditure (% GDP)	-0.1315	0.2285	0.1652	0.1458	0.0750	-0.2124
	(0.1876)	(0.2196)	(0.1313)	(0.1022)	(0.0791)	(0.1793)
Gov.Expenditure (% GDP) (-1)	-0.0238	-0.1794	-0.0303	0.0655	0.1091	-0.0253
	(0.2011)	(0.2338)	(0.1406)	(0.1156)	(0.0895)	(0.1961)
Implied Long Run Coefficients						
Inflation	**-9.2948**	3.0208	**5.6276**	**6.2993**	**4.6286**	**-8.9190**
F-statistic	(24.2019)	(2.5204)	(18.9882)	(38.7763)	(45.0227)	(22.9826)
Probability	(0.0001)	(0.1267)	(0.0003)	(0.0000)	(0.0000)	(0.0001)
Squared inflation	**59.0424**	0.8762	**-34.2377**	**-41.4919**	**-32.5434**	**54.2401**
F-statistic	(9.8660)	(0.0022)	(7.0757)	(17.0771)	(23.0390)	(8.5741)
Probability	(0.0047)	(0.9631)	(0.0143)	(0.0004)	(0.0001)	(0.0078)
Growth	-10.2202	8.1448	*9.2243*	**9.9770**	2.3044	-12.5533
F-statistic	(1.6483)	(0.9875)	(3.0792)	(5.1049)	(0.5873)	(2.6633)
Probability	(0.2126)	(0.3312)	(0.0932)	(0.0341)	(0.4516)	(0.1169)
Gov. Expenditure (%GDP)	-0.2402	0.0683	0.2009	0.3172	0.2475	-0.3749
F-statistic	(0.3684)	(0.0255)	(0.5705)	(2.2663)	(2.7141)	(0.9549)
Probability	(0.5501)	(0.8745)	(0.4581)	(0.1464)	(0.1137)	(0.3391)
Inflation threshold	**7.87%**	-172.38%	**8.22%**	**7.59%**	**7.11%**	**8.22%**
F-statistic	(62.4059)	(0.0024)	(36.6652)	(109.9573)	(186.5355)	(48.7659)
Probability	(0.0000)	(0.9617)	(0.0000)	(0.0000)	(0.0000)	(0.0000)

(*) In each equation the dependent variable is log(y/(1-y)) where y is the Gini coefficient, or the lowest, second, third, fourth and highest quintile income share respectively.

Bold = significant at 5%. Italic = significant at 10%.

Table 2C - Static regressions for the US - Linear model(*)						
Estimation results	**GINI**	**lowest**	**second**	**third**	**fourth**	**highest**
Constant	**0.9403**	**-3.7877**	**-3.36746**	**-2.5632**	**-1.3378**	**0.9916**
	(0.2507)	(0.1217)	(0.2191)	(0.1664)	(0.0723)	(0.2167)
Inflation	**-4.6701**	**3.5773**	**3.57908**	**3.1378**	**1.4863**	**-4.6302**
	(0.4855)	(0.2357)	(0.4243)	(0.3222)	(0.1400)	(0.4196)
Real GDP growth	**-34.1315**	**14.0393**	**32.37587**	**25.5039**	*4.0983*	**-31.0482**
	(7.8481)	(3.8107)	(6.8590)	(5.2087)	(2.2639)	(6.7827)
Diagnostics						
Adjusted R-squared	0.7401	0.8876	0.6841	0.7438	0.8020	0.7918
F-statistic	46.5694	127.3565	35.6531	47.4421	65.8019	61.8311
Prob(F-statistic)	0.0000	0.0000	0.0000	0.0000	0.0000	0.0000

(*) Dependent variable is $\log(y/(1-y))$ where y is the Gini coefficient, or the lowest, second, third, fourth and highest quintile income share respectively.

Bold = significant at 5%. Italic = significant at 10%.

Table 3C – Dynamic regressions for the US - Linear model (*)						
Estimation results	GINI	lowest	second	third	fourth	highest
Constant	*0.2224*	**-2.3076**	**-0.5774**	**-0.4057**	**-0.3607**	0.2152
	(0.1254)	(0.6663)	(0.2536)	(0.1911)	(0.1592)	(0.1279)
Lagged dependent variable	**0.8105**	*0.3623*	**0.8245**	**0.8505**	**0.7545**	**0.8105**
	(0.0926)	(0.1814)	(0.0826)	(0.0812)	(0.1124)	(0.0955)
Inflation	-1.5184	**4.7385**	0.9766	0.8679	0.2832	-1.6676
	(1.4100)	(1.4030)	(1.1048)	(0.8215)	(0.5200)	(1.2818)
Inflation(-1)	0.5928	**-2.6498**	-0.3780	-0.3597	0.2002	0.7230
	(1.3059)	(1.2875)	(1.0176)	(0.7555)	(0.5900)	(1.2061)
Real GDP growth	11.5714	0.8371	-9.2678	-4.7667	-4.4288	9.2704
	(8.5015)	(8.9241)	(6.2615)	(4.8765)	(3.8328)	(8.1563)
Real GDP growth(-1)	**-19.1759**	4.8410	**14.3156**	*9.0742*	6.2355	*-15.7091*
	(8.2837)	(8.4338)	(6.0836)	(4.7891)	(3.6795)	(7.9430)
Adjusted R-squared	0.9613	0.9242	0.9655	0.9710	0.9297	0.9620
F-statistic	155.0752	76.6021	174.4455	208.8571	83.0500	157.8235
Prob(F-statistic)	0.0000	0.0000	0.0000	0.0000	0.0000	0.0000
Implied Long Run Coefficients						
Inflation	**-4.8839**	**3.2754**	**3.4116**	**3.4002**	**1.9695**	**-4.9837**
F-statistic	(16.2940)	(88.3424)	(13.6957)	(13.8100)	(18.8112)	(18.1193)
Probability	(0.0004)	(0.0000)	(0.0010)	(0.0010)	(0.0002)	(0.0002)
Growth	*-40.1225*	8.9042	*28.7684*	*28.8170*	7.3603	-33.9716
F-statistic	(3.1241)	(2.0915)	(2.9886)	(3.0787)	(1.0853)	(2.7347)
Probability	(0.0889)	(0.1601)	(0.0957)	(0.0911)	(0.3071)	(0.1102)

(*) Dependent variable is log(y/(1-y)) where y is the Gini coefficient, or the lowest, second, third, fourth and highest quintile income share respectively.

Bold = significant at 5%. Italic = significant at 10%.

Estimation results	GINI	lowest	second	third	fourth	highest
Table 2D - Static regressions for the US using the post-1983 dummy (*)						
Constant	0.2158028	**-3.57378**	**-2.67293**	**-2.059717**	**-1.286791**	**0.3957129**
	(0.1707)	(0.1354)	(0.1362)	(0.1035)	(0.0575)	(0.1519)
Inflation	**-12.24377**	**4.285587**	**8.83105**	**8.1306755**	**4.836299**	**-11.71381**
	(1.5585)	(1.2361)	(1.2436)	(0.9446)	(0.5245)	(1.3864)
Squared Inflation	**98.319752**	-14.6139	**-75.30243**	**-65.77626**	**-33.50033**	**88.929133**
	(15.6605)	(12.4211)	(12.4961)	(9.4914)	(5.2699)	(13.9315)
Growth	-8.91546	*7.793679*	**9.78007**	**8.1916891**	0.110181	*-9.636591*
	(5.3675)	(4.2572)	(4.2829)	(3.2531)	(1.8062)	(4.7749)
Dummy 1983-99	**0.104579**	**-0.03405**	**-0.10442**	**-0.073243**	-0.00152	**0.0842465**
	(0.0157)	(0.0124)	(0.0125)	(0.0095)	(0.0053)	(0.0140)
Inflation threshold	**6.23%**	*14.66%*	**5.86%**	**6.18%**	**7.22%**	**6.59%**
F-statistic	(377.7757)	(2.9761)	(395.8198)	(464.5610)	(264.5079)	(339.0871)
Probability	(0.0000)	(0.0955)	(0.0000)	(0.0000)	(0.0000)	(0.0000)
Adjusted R-squared	0.9182861	0.905701	0.917205	0.9328107	0.915267	0.9306208
F-statistic	90.902561	77.83665	89.62509	112.06653	87.41441	108.30828
Prob(F-statistic)	0.0000	0.0000	0.0000	0.0000	0.0000	0.0000

(*) In each equation the dependent variable is log(y/(1-y)) where y is the Gini coefficient, or the lowest, second, third, fourth and highest quintile income share respectively.

Bold = significant at 5%. Italic = significant at 10%.

Table 4 – Gini coefficients in 15 OECD countries															
income net, household, selected by country-specific homogeneous source															
Year	AUS	BEL	CAN	DNK	FIN	FRA	DEW	IRL	ITA	NLD	NOR	PRT	SWE	GBR	USA
1960							38.0								
1961															
1962											37.5				
1963															
1964							38.0								
1965															
1966															
1967									40.0		36.0				
1968							38.7		40.0						
1969									40.0					26.3	
1970							39.2		38.0						
1971									40.0						
1972									38.0						
1973							37.0	36.7	42.0		37.5				
1974									41.0					28.6	34.0
1975			31.4				36.6		39.0						
1976									35.0		37.3		33.1		
1977									36.0						
1978							36.4		36.0						
1979						31.7			37.0		31.2				32.6
1980							36.6	36.0	37.0			36.8	32.4		
1981			30.5			27.9			32.0				32.5		
1982									31.0				30.7		
1983						31.9	33.9								
1984							33.4				30.6		31.8		
1985		26.2					35.2			31.4			31.2		
1986	41.0								33.0		33.1		31.7	33.3	35.2
1987			31.2	31.3	26.2			35.2	34.4				31.7		
1988		26.6											32.2		
1989									33.9				31.3		
1990	42.0											36.8	32.5		
1991			31.5		26.1				32.5	31.8	33.3				35.2
1992		26.9		30.0						31.7			32.4		
1993									36.3	32.0					
1994										32.4					
1995	44.0								36.2						
1996	44.0														
obs 60-96	4	3	4	2	2	3	11	3	22	4	9	2	12	3	4
obs 73-96	4	3	4	2	2	3	7	3	16	4	7	2	12	2	4
adjusted*	3	3	4	2	2	3	6	3	10	2	6	2	8	2	4

Tot 88   *Tot 75*   *Tot 60*

Observations in shaded cells are excluded from the benchmark sample.

Table 5 - Annual CPI inflation (%) in selected OECD countries, 1973-1996					
Country	Mean	Median	Maximum	Minimum	Std. Dev.
AUS	7.87	8.23	15.31	0.99	4.09
BEL	5.18	4.47	12.76	1.16	3.40
CAN	6.07	5.30	12.42	0.16	3.49
DEW(*)	3.68	3.56	7.04	-0.22	2.16
DNK	6.57	5.53	15.29	1.27	4.07
FIN	7.37	6.83	17.84	0.63	4.88
FRA	6.76	6.60	13.66	1.66	4.32
GBR	8.62	6.93	24.21	1.56	5.96
IRL	8.85	6.54	20.63	1.40	6.76
ITA	10.65	9.99	21.08	4.01	5.72
NLD	4.09	3.18	10.22	-0.69	3.02
NOR	6.68	6.94	13.65	1.25	3.51
PRT	15.95	14.26	33.06	3.09	8.18
SWE	7.44	7.69	13.69	0.80	3.40
USA	5.71	4.56	13.54	1.90	3.21
All countries	7.43	6.33	33.06	-0.69	5.43

(*) West Germany: 1973-1990.

Source: OECD Main Economic Indicators

Table 6 - Annual real GDP growth (%) in selected OECD countries, 1973-1996					
Country	Mean	Median	Maximum	Minimum	Std. Dev.
AUS	2.91	2.98	6.60	-2.76	2.33
BEL	2.03	2.17	6.04	-2.49	2.27
CAN	3.01	3.69	7.57	-4.57	2.91
DEW(*)	2.19	2.46	5.56	-1.52	2.24
DNK	1.90	1.69	6.73	-1.70	2.17
FIN	2.14	2.91	7.92	-9.60	3.93
FRA	2.09	2.41	5.56	-1.59	1.74
GBR	2.07	2.46	7.07	-2.64	2.62
IRL	4.33	3.72	10.20	-1.09	2.91
ITA	2.45	2.70	7.49	-4.00	2.63
NLD	2.25	2.48	5.04	-1.45	1.71
NOR	3.46	3.83	6.96	-0.41	2.00
PRT	3.67	2.97	12.39	-2.77	3.66
SWE	1.54	1.92	4.08	-2.30	2.04
USA	2.42	2.84	6.99	-3.02	2.55
All countries	2.57	2.67	12.39	-9.60	2.65

(*) West Germany: 1973-1990.

Source: OECD Main Economic Indicators

| Table 7 – Alternative model estimates for the OECD countries				
Dependent Variable: Gini index				
Method: fixed effects estimator				
Sample: 1973 1996				
Total panel (unbalanced) observations: 60				
Number of cross-sections used: 15				
White Heteroskedasticity-Consistent Standard Errors & Covariance				
Variable	**Model 1**	**Model 2**	**Model 3**	**Model 4**
Inflation	**-0.8981**	-0.1369	**-0.845**	**-1.0874**
	(0.1997)	(0.0932)	(0.2429)	(0.1989)
Squared inflation	**3.5280**		**3.3970**	**3.5567**
	(0.7051)		(0.8480)	(0.6669)
Real GDP growth	**2.0704**	**1.3636**	**1.9984**	**1.8227**
	(0.5425)	(0.5068)	(0.5746)	(0.4915)
Gov. Exp. (%GDP)			**-0.6891**	
			(0.2598)	
Dummy 73-81				**0.0179**
				(0.0063)
Adjusted R-squared	0.8192	0.7937	0.7984	0.8422
Inflation threshold	**12.73%**		**12.44%**	**15.29%**
F-Statistic	213.3102		175.2571	141.4988

Estimated fixed effects not reported in table for space reasons.
Standard errors in parenthesis.
Bold = significant at 5%.

Table 8	
Gini sacrifice ratios	
Inflation	**Elasticity**
1%	-0.8276
2%	-0.7570
3%	-0.6864
4%	-0.6159
5%	-0.5453
6%	-0.4748
7%	-0.4042
8%	-0.3336
9%	-0.2631
10%	-0.1925
11%	-0.1220
12%	-0.0514
13%	0.0192
14%	0.0897
15%	0.1603
16%	0.2308
17%	0.3014
18%	0.3720
19%	0.4425
20%	0.5131

Table 9						
Hypothetical shares of total income by quintiles						**Gini coefficient**
	q20	**q40**	**q60**	**q80**	**q100**	
case 1	5	12	18	25	40	33.2
case 2	5.5	12	18	25	39.5	32.4

Table 10 - Sensitivity of estimated coefficients to country exclusion				
Excluded country	**Inflation**	**Squared inflation**	**Growth**	**Inflation threshold**
Australia	**-0.8765**	**3.4426**	**2.0527**	12.73%
	(0.2449)	(0.8696)	(0.5696)	
Belgium	**-0.9390**	**3.6928**	**2.1264**	12.71%
	(0.2129)	(0.7542)	(0.5585)	
Canada	**-0.8999**	**3.5315**	**2.2986**	12.74%
	(0.2105)	(0.7375)	(0.5906)	
West Germany	**-1.0321**	**4.0712**	**2.2324**	12.68%
	(0.2071)	(0.7247)	(0.5804)	
Denmark	**-0.9738**	**3.8374**	**2.1464**	12.69%
	(0.2030)	(0.7115)	(0.5476)	
Finland	**-0.9053**	**3.5499**	**2.1503**	12.75%
	(0.2020)	(0.7124)	(0.5658)	
France	**-0.8814**	**3.4725**	**2.0403**	12.69%
	(0.1996)	(0.7045)	(0.5435)	
Great Britain	**-0.8306**	**3.3198**	**1.9379**	12.51%
	(0.2197)	(0.7611)	(0.5820)	
Ireland	**-1.0642**	**4.0027**	**2.4301**	13.29%
	(0.2015)	(0.7496)	(0.5585)	
Italy	**-0.7080**	**2.7898**	**1.6003**	12.69%
	(0.1828)	(0.6835)	(0.4861)	
Netherlands	**-0.9004**	**3.5372**	**2.0738**	12.73%
	(0.2000)	(0.7062)	(0.5432)	
Norway	**-0.7654**	**3.0240**	**1.6733**	12.65%
	(0.2011)	(0.7111)	(0.5887)	
Portugal	**-0.8920**	**3.4893**	**2.0721**	12.78%
	(0.2137)	(1.2638)	(0.5695)	
Sweden	**-0.9076**	**3.5587**	**2.1391**	12.75%
	(0.2069)	(0.7194)	(0.5544)	
USA	**-0.8857**	**3.4817**	**2.0497**	12.72%
	(0.2339)	(0.8251)	(0.5730)	
All countries	**-0.8981**	**3.5280**	**2.0704**	12.73%
	(0.1997)	(0.7051)	(0.5425)	

Standard errors in parenthesis.
Bold = significant at 5%.

Figure 1 – Short-run and Long-run Effects of Restrictive Monetary Policy on Income Distribution

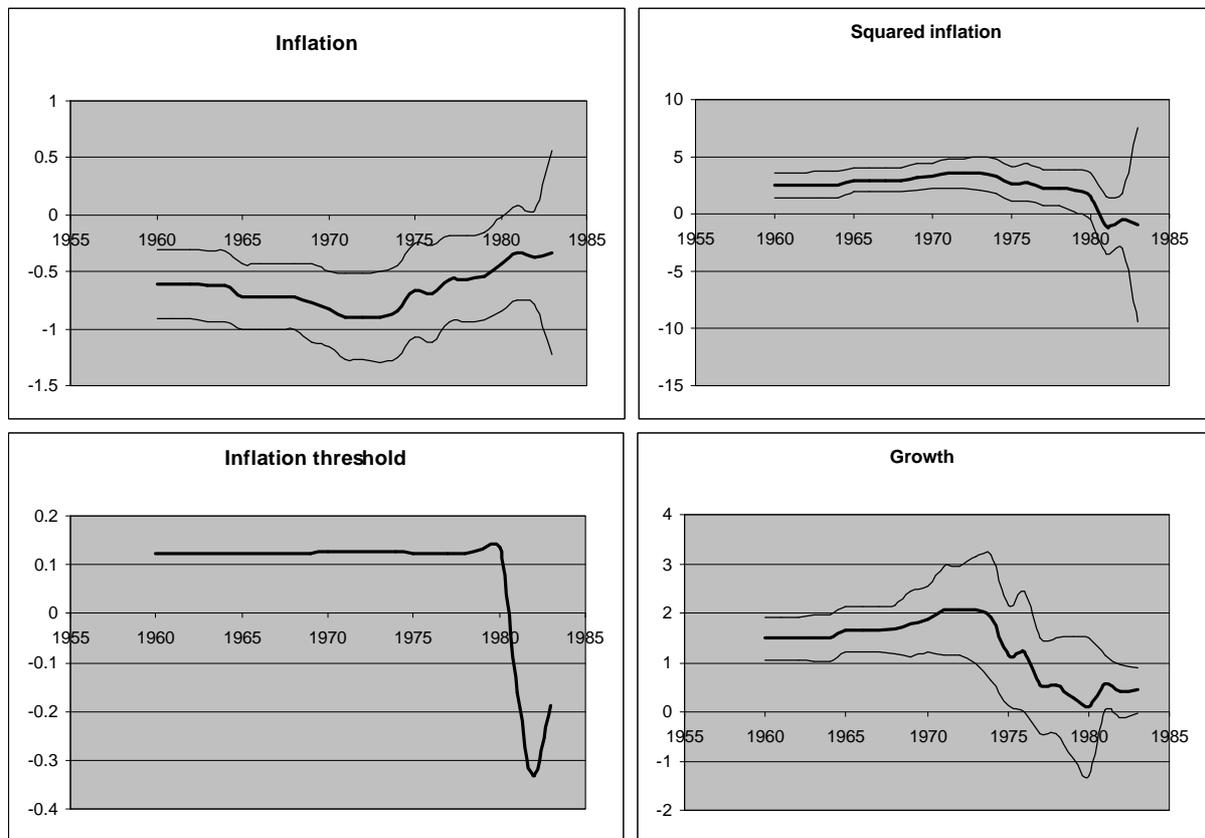* Channels activated only or mainly in high inflation economies.
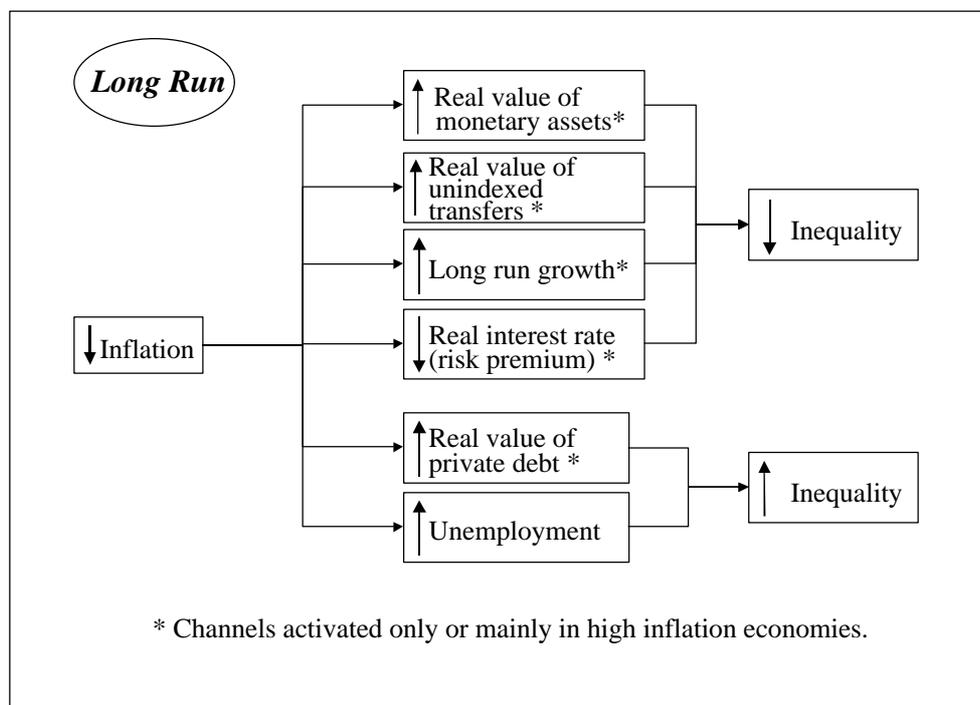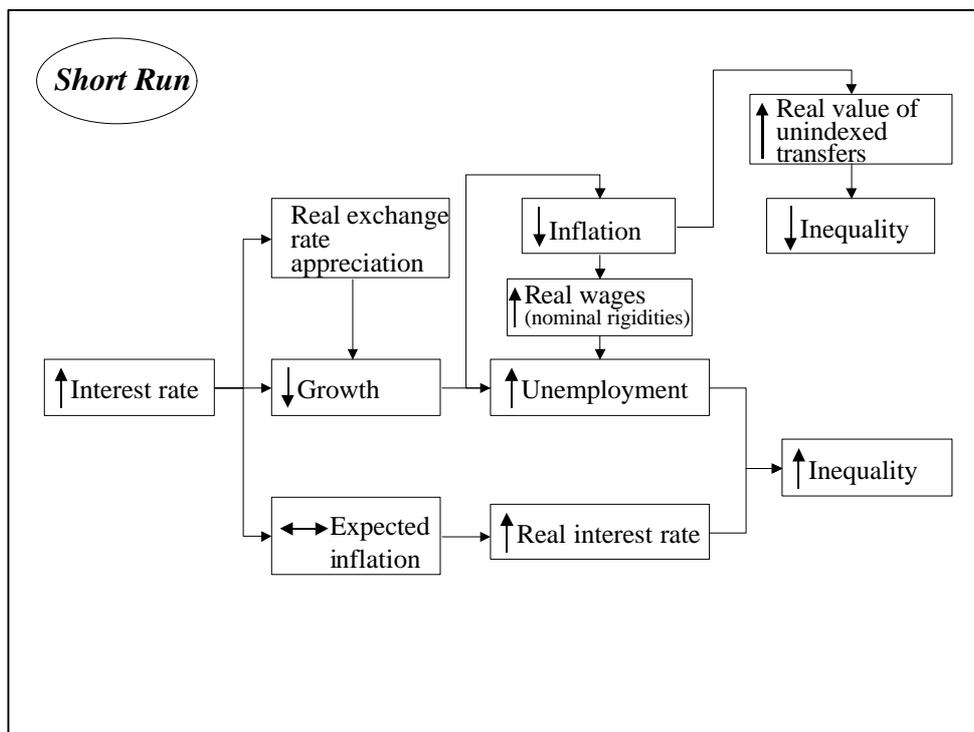
Figure 2 – Gini index and quintiles income shares, USA, 1966-1999
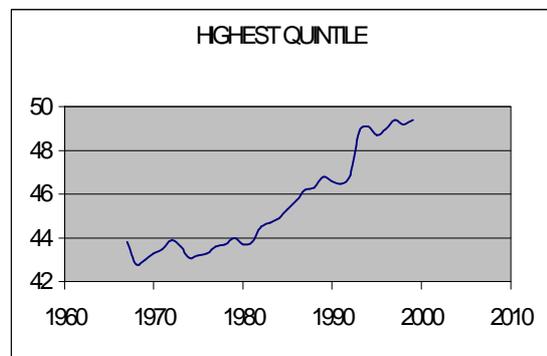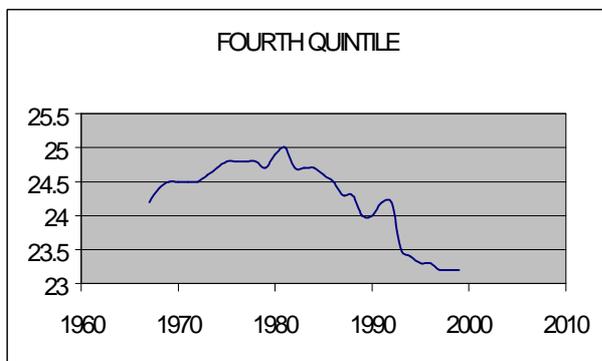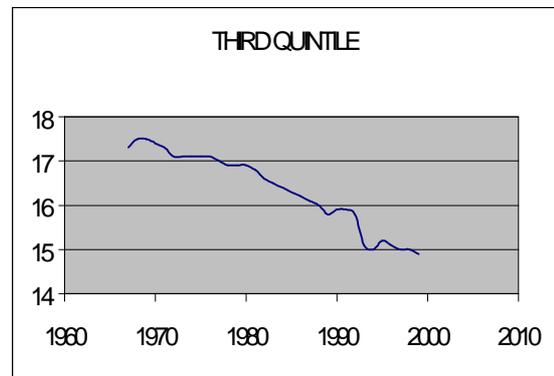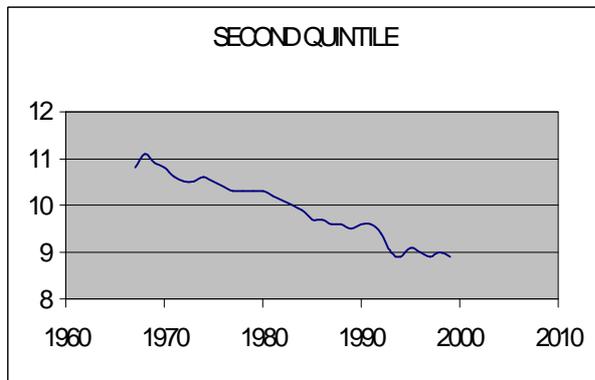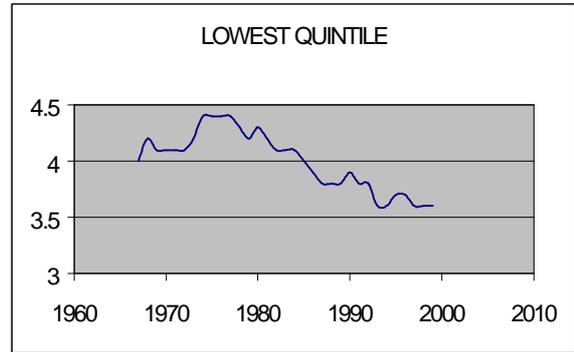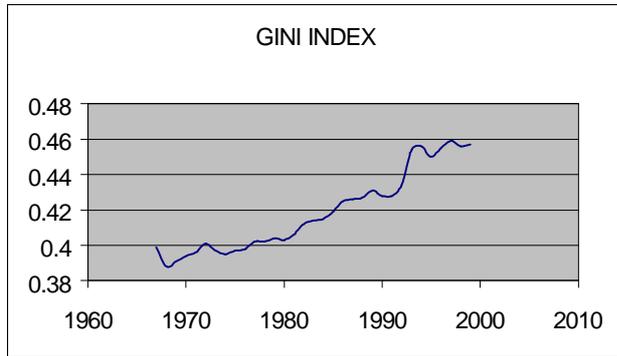
Figure 3 – Coefficient stability in the Gini static model for the US
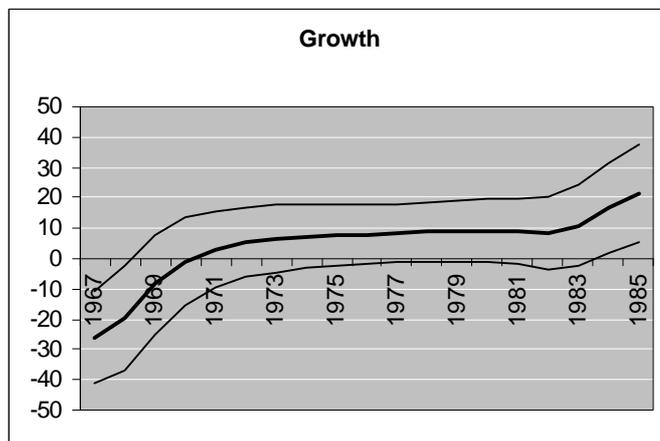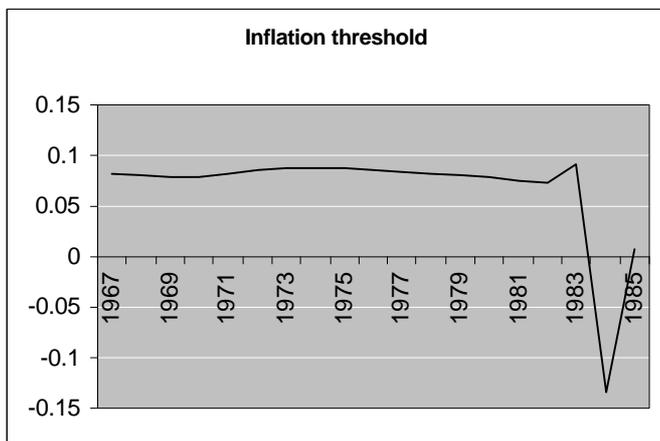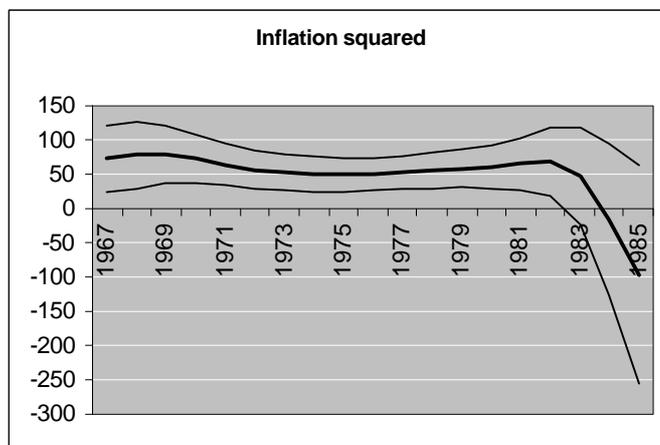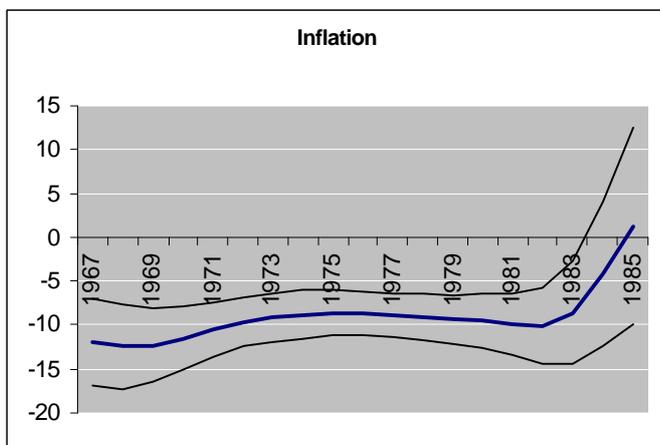
Figure 4 – Coefficient stability in the Gini model for the OECD countries

Contents

Tables and figures