Inflation and Factor Shares*

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Abstract

We use results from the literature on the determinants of price-cost margins to derive an equation relating labor's share of national income to the inflation rate (as well as to the output gap, the unemployment rate and the capital stock per worker). The equation is tested with a panel of 15 OECD countries. We obtain a robust positive relationship between inflation and the labor share. Our results suggest that disinflation is not distributively neutral, provide empirical support for the distinct concern about price stability shown by trade unions and employers' organizations, and help explaining the negative impact of inflation on growth.

Keywords: Inflation, Functional Distribution of Income, Markups.
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1 Introduction

Disinflation during the eighties and the nineties has been accompanied by a significant rise in the profit share of national income in most OECD countries (Blanchard, 1997) or, equivalently, by a reduction in the labor share (see Figure 1). This suggests the possibility that disinflation -or more generally, changes in the level of inflation- could be non-neutral with respect to the distribution of income. In fact, low-inflation oriented central banks have traditionally been characterized as "conservative". Furthermore, the distinct concern about inflation shown by employers' organizations in comparison to the one exhibited by trade unions can be rationalized only if one assumes that inflation has a different impact on their income. The aim of this paper is to empirically assess the possible effect of inflation on factor shares.

The literature on the distributive impact of inflation has typically focused on personal income inequality. The issue has been investigated by using information on the quintiles distribution of familiar income, or either by means of some measure of income dispersion like the Gini coefficient, and the results have been somehow inconclusive.[1] By contrast, the analysis of the impact of inflation on the functional distribution of income has received almost no academic attention, despite its great importance. Basically, this importance derives from the fact that social agents intervening in the political economy of monetary policy and the fight against inflation -like trade unions and employers' organizations- are mostly defined in terms of their type of income. Hence from a positive political economy point of view, analyzing the impact of inflation on the factor shares is a key issue. On the other hand, the means by which inflation may affect personal income inequality can be due to a large extent to the fact that different types of income are differently affected by inflation, and that the composition of income (between wages, profits, interests and others) is very different across individuals and systematically related to total income. Thus the consequences of inflation upon inequality may largely be the indirect result of the effects of inflation upon factor shares.

In the following we study the determinants of the labor share in a panel of 15 OECD countries with data from 1960 to 1999. Our results indicate that

[1] See Blinder and Esaki (1978), Jantti (1994), Sarel (1997) and Bulir (1998). They respectively find that the distributive impact of inflation is weak progressive, significant progressive, non-significant, and regressive.

inflation is a robust positively signed determinant of the labor share. This positive effect comes up in pooled estimations as well as in almost all country-specific estimations.

This result suggest that the eventual benefits of low inflation are not equally shared by all production factors, at least in the short and medium run. As a consequence, it provides new foundations for the inflationary bias that -according to some macroeconomic models- trade unions may cause, and makes the case for going further than the representative agent model to analyze the overall consequences of monetary policy and disinflation. On the other hand, it helps explaining the negative impact of inflation on capital growth. The literature on this topic has pointed at two channels of influence to produce this negative effect: the *efficiency channel* by which inflation would lower total factor productivity; and the *accumulation effect* by which inflation would lower investment, at least temporarily (see Andrés and Hernando,1999, and the references therein). The reduction that inflation operates in the capital share -according to our results- may explain this negative effect on investment.

The empirical study is framed into a simple theoretical model that builds to a large extent on the relationship between inflation and market power thoroughly investigated by the staggered pricing literature. In the next section we lay out the arguments. In Section 3 we proceed to the empirical analysis. In the last Section we summarize and offer some concluding comments.

2 Theoretical Framework

Let Y be output, L labor, $k \equiv K/L$ capital per worker, p output price, w nominal wage, and μ the markup over marginal costs. We assume a degree-one homogeneous production function, and a monopolistic competition economy implying a markup-over-marginal-costs pricing equation:

$$Y = L \cdot F(k), \tag{1}$$

$$p = \frac{w}{\partial Y/\partial L} \cdot \mu(.); \tag{2}$$

From these two equations we obtain

$$s \equiv \frac{\omega L}{Y} = \frac{1}{\mu(.)} \cdot \Phi(k). \tag{3}$$

where ω is real wage ($\omega = w/p$), and $\Phi(k) \equiv 1 - k\left[F'(k)/F(k)\right]$. Hence, together with technology, variables affecting firms' margins determine the relative distribution of net output between capital and labor.[2]

The literature on price determination shows that markups may depend on a series of factors like industry concentration, collusion, demand elasticity, and potential entry. Most changes on these variables are likely to show a uncorrelated idiosyncratic profile across industries. Nonetheless, they may simultaneously be affected by the business cycle. Hence from the macroeconomic perspective, the business cycle is a potentially important determinant of average margins, though their overall cyclical behavior is still under discussion (see Rotemberg and Woodford, 1999). Furthermore, to the extent that agreements between unions and employers are efficient -i.e., they simultaneously set both wages and employment- markups also depend on unions' bargaining power and the elasticity of substitution between wage and employment in unions' preferences (Dowrick, 1990; see also Kalecki, 1954). Since higher unemployment is likely to decrease both union's bargaining power and the substitutability of employment by wages, it may also constitute a significant macroeconomic determinant of markups. In particular, we may expect a positive relationship between markups and the unemployment rate.[3]

Besides those variables usually considered as determinants of margins, we are most interested in the series of arguments suggesting that inflation may affect markups and real profits. First, fixed costs of price adjustments imply that staggered pricing following (S, s) rules are optimal; where S and s are relative prices of output with respect to marginal costs (Sheshinsky and Weiss, 1977 and 1983). With higher steady inflation, both limits S and s tend to increase. In the case of monopolistic firms, this will result in higher or in lower average markups depending on the form of the profit and demand functions, and the discounting rate (Konieczny, 1990, and Benabou and Konieczny, 1994).

Second, with more than one firm, steady inflation will tend to increase com-

[2]Kalecki (1938) and (1954) was probably the first to propose a theory of the aggregate income distribution hinging on the determinants of firms' market power.

[3]The relevant unemployment rate for markups at time t is likely to be the one corresponding to time $t-1$, since wages are usually negotiated one period in advance. The same would occur if the relevant unemployment rate is the one expected for the next period, but adaptive expectations are used. In any case, due to the strong inertia shown by the unemployment rate in all countries, our empirical results are almost identical whatever we use lagged or contemporaneous unemployment rates.

petition. The higher the inflation rate, the wider will be the S-s spread. Hence, with uncoordinated price increases, higher inflation gives rise to a wider price dispersion that offers room for more intense consumer search. This in turn creates incentives for more competition that tends to lower down margins over marginal costs. Moreover, markups in the lower segments of every firm S-s cycle receive higher weights -in terms of sales- when computing average margins (Benabou, 1988 and 1992a).[4] At the empirical level, Benabou (1992b) obtained evidence of a negative relationship between markups and the level of inflation, as well as on the non-significance of inflation uncertainty, using data from the USA retail trade sector.

Furthermore, some other elements can add to a negative effect of inflation upon profits per unit of output. First to the extent that markup pricing by firms is not always applied on replacement or expected costs, but on historical costs,[5] inflation reduces firms' capitalization and real profits. Second, higher inflation rises several costs such as the menu ones. Though these cost increases could reduce the marginal productivity of labor -and therefore, real wages- in the same proportion as profits, so that factor shares would be unaltered, inasmuch as they take the form of higher fixed costs they will tend reduce profits relatively to wages (at least in the short and medium run). Third, sluggish price adjustment due to menu costs implies a higher average distance with respect to the optimal price, thus reducing profits. And finally, the non-neutralities with respect to inflation built in most countries' tax systems biases capital allocation towards less profitable, but fiscally less exposed, investments.

According to all these arguments we consider the following equation on the markup determinants:

$$\mu = \mu\left(\pi, y, u_{-\tau}\right); \tag{4}$$

[4] In opposition to this result, it has been argued that inflation variability may reduce the value of information that is obtained by search, therefore reducing it and increasing firms' market power (see, for instance, Tommasi, 1994). Thus, to the extent that higher inflation tends to imply higher uncertainty about the price level, there would be a countervailing force against the negative relationship between inflation and margins. Nonetheless, Benabou and Gertner (1993) show that the sign of this uncertainty effect is ambiguous, and that it depends on the possibility of acquiring more information.

[5] Valuation of inputs on a historical basis has been the only method legally admitted in many countries in order to fulfill legal accounting obligations, and to compute costs and profits. Some portion of firms may then be induced to use this misleading information on costs to set prices.

where π is the inflation rate, y is the output gap, and $u_{-\tau}$ is (τ) lagged unemployment. Thus (3) can be rewritten as:

$$s = \frac{1}{\mu\left(\pi, y, u_{-\tau}\right)} \cdot \Phi(k) \tag{5}$$

Now, log-linearizing this equation we have:

$$\ln s = \gamma_0 + \gamma_1 \ \ln \pi + \gamma_2 \ \ln y + \gamma_3 \ \ln u_{-\tau} + \gamma_4 \ \ln k \tag{6}$$

Our previous discussion imply a likely positive sign for the impact of inflation on the labor share, and a negative sign for the elasticity on unemployment. On the other hand, the literature on the cyclical behavior of markups does not provide definitive results for γ_2. Though the usual correlation obtained between contemporaneous output gap and the labor share has been negative (see Rotemberg and Woodford, 1999, and the references therein). Finally there are no theoretical restrictions on the sign of the derivative of $\Phi(k)$, and therefore on γ_4.

A side comment on the likely (very) long run effects of inflation may be convenient. Inasmuch as $\partial\mu/\partial\pi$ is negative, higher inflation will tend to bring about lower profits. In the case of monopolistic competition industries with no barriers to entry -i.e., with no monopoly power in the long run- an average inflation higher than the expected at the time of carrying out fixed investments may give rise to persistent losses and, therefore, to a reduction in the long run number of firms. The smaller number of firms may in turn induce less consumer search (Diamond, 1993) and may sustain in the long run an equilibrium with higher markups. This, together with the reduction in the inefficiencies resulting from monopolistic competition with increasing returns to scale (Benabou, 1987 and 1993) may -at least, to some extent- offset the initial rise in the labor share caused by inflation. Hence the impact on the labor share of a permanent increase in the average level of inflation, may lower in the long run. This would be consistent with the negative effect of inflation on investment and growth being only temporary, as suggested by the inflation-and-growth literature. Symmetrically, the rise in profits prompted by disinflation may last for a fairly long time but may be not permanent, since it would likely be followed by an increase in industries' capacity which would eventually work against the markup increase. This hypothesis could deserve some specific empirical work in the future.

3 Empirical Analysis

The labor share was computed in two different ways. The first series was obtained as the ratio of employees' compensation in the business sector to the value added in the business sector at factor prices. In the second choice we added to the employees' compensation, the labor income imputed to the self-employed individuals, obtained by multiplying their number by the average wage in the business sector. Both series give rise to very similar estimations. The results we report correspond to the first definition of the labor share. Inflation is computed using the Consumer Price Index; i.e., $1 + \pi_{it} \equiv CPI_{it}/CPI_{it-1}$. The output gap y_{it} is defined as $Y_{it}/\overline{Y}_{it}$, where Y_{it} is country i's total output at time t and $\overline{Y}_{it}$ is Hodrick-Prescott detrended output (with the usual smoothing parameter).

The main data source is the OECD Economic Outlook data diskette. CPI data are from the Main Economic Indicators. The capital stock is from De la Fuente and Doménech (2000) which is obtained from the OECD data on investment, enlarged from below by using other complementary sources, and correcting for differences in PPP. Oil prices -which are used only as an instrument- are from IMF Financial Statistics.

We carried out two blocks of estimations. In the first block we pool the 15 OECD countries for which we have an almost complete time series during the period 1960-99,[6]. In the second block we perform country-specific estimates of the same equation to check the robustness of our results.

3.1 Pooled Estimations

As noted by Judson and Owen (1996), the fixed effects model is generally more appropriate than the random effects model for many macro dataset. There are two reasons. First, the individual effects are likely to be correlated with the other regressors whenever they are substituting for omitted variables. And second, the selected group of countries considered in this paper -as usually occurs- is far from being a random sample of countries. From (6), the fixed effects model to be

[6]We limited our study to those OECD countries for which we had at least a 25 years complete time series set. This left us with 16 countries. From these we excluded Norway since many economic variables -such as the labor share- have a very idiosyncratic path due to the very erratic value of the oil and natural gas production which accounts up to the 20% of the business sector output in some years.

estimated is:

$$\ln s_{it} = \gamma_{0i} + \gamma_1 \ln(1 + \pi_{it}) + \gamma_2 \ln u_{it-1} + \gamma_3 \ln y_{it} + \gamma_4 \ln k_{it} + \varepsilon_{it} \qquad (7)$$

where γ_{0i} is country i's fixed-effect, and ε_{it} is the random disturbance.

In Table I we report the results for equation (7) estimations with different methods. The first column correspond to the Least Squares Dummy Variable estimator (LSDV). These estimations are obtained under the assumption that errors are contemporaneously uncorrelated and cross-section -as well as time-series- homoskedastic. The second column correspond to the results obtained with the Seemingly Unrelated Equations estimator (SUR); i.e., assuming that residuals are both cross-section heteroskedastic and contemporaneously correlated. As it can be observed, all the explanatory variables are statistically very significant in both estimations, and have the expected signs whenever we had any hypothesis at all (i.e., for all variables except k). Moreover, coefficient estimates are fairly similar for the two methods. To get a grasp on the quantitative importance of the positive effect of inflation, note that with an estimated coefficient of 0.71 and a labor share of about 65% of the national income, going from an average inflation of 5% to an average inflation of 2% would reduce the labor share in about 1.3 percent points.

Simultaneous determination of the explained variable and regressors could cause some endogeneity problems to our estimations. In order to evaluate the quantitative effect of this potential problem we carry out two GMM estimations with different sets of instruments. In principle, the use of instrumental variables can deal with the endogeneity problem, Although the GMM is sometimes subject to skepticism since in practice it is difficult to find appropriate instruments, these estimations may provide an additional test on the robustness of our results.

Taking the first difference of (7) to eliminate γ_{0i}, we have

$$\Delta \ln s_{it} = \gamma_1 \Delta \ln(1 + \pi_{it}) + \gamma_2 \Delta \ln u_{it-1} + \gamma_3 \Delta \ln y_{it} + \gamma_4 \Delta \ln k_{it} + \Delta \varepsilon_{it} \qquad (8)$$

The natural instruments are lags of regressors. Assuming that ε_{it} is serially uncorrelated, lags of regressors dated $t-2$ and earlier are valid instruments (except for the output gap since, by construction, the Hodrick-Prescott filter implies that we should use an unreasonable number of lags to avoid any correlation between errors and instruments). We use two different sets of predetermined instruments: IV_1 and IV_2. The set IV_1 includes two-periods lagged unemployment growth rate,

inflation rate, capital-labor ratio, and GDP growth rate, as well as the growth rate of the oil nominal price lagged two and three periods. IV_2 includes two-periods lags of the level as well as of the first difference of the unemployment rate, the inflation rate and the capital-labor ratio; and also the nominal price of oil lagged two and three periods.[7] Results are reported in the third and fourth columns of Table 1. As it can be observed, signs and statistical significance for the inflation and the output gap are the same; and the quantitative values of their coefficients change only slightly. The signs of the two other variables are also the same, but the unemployment rate becomes not significant with the second set of instruments; and the capital-labor ratio is not significant with either of the instrument sets.

The differences between the GMM results and those previously obtained by the SUR method can be statistically compared by means of the Hausman-Wu test. The null hypothesis is that the regressors in equation (7) are exogenous. Probabilities for this test are 0.097 and 0.098 for IV_1 and IV_2, respectively. Hence we do not reject the null hypothesis in any of the two cases. Therefore, the potential endogeneity problem does not seem to cause any significant bias in the coefficient estimates using the SUR method.

In order to check for the robustness of the effect of inflation under different specifications, we also performed estimations alternatively eliminating the rest of explanatory variables. Results are reported in Table 2. As it can be observed, inflation was always found to be positive and very significant, and the coefficients obtained were quite similar to those in Table 1. Hence the estimated impact of inflation does not seem to depend on the choice of any particular set of additional explanatory variables.

3.2 Country-Specific Estimations

Figure 1 suggests that the inflation rate, besides being a significant variable in pooled estimates of the labor share, can also be an important factor explaining the evolution of the labor share in every of most OECD countries. We now estimate equation (6) for each country, yet allowing in the estimation procedure for the error terms to be heteroskedastic and contemporaneously correlated between

[7] These two sets of instruments work statistically well since we do not reject the overidentifying restrictions in either case. The p-values for the Sargan test are, respectively, 0.07 and 0.14.

countries (i.e., we use the SUR estimation method). In this way we take account of the common shocks to all countries -as the oil ones- that occurred during the period encompassed by our data. The equation to be estimated is:

$$\ln s_{it} = \gamma_{0i} + \gamma_{1i} \ln (1 + \pi_{it}) + \gamma_{2i} \ln u_{it-1} + \gamma_{3i} \ln y_{it} + \gamma_{4i} \ln k_{it} + \varepsilon_{it}. \qquad (9)$$

Figure 2 suggests that equation (9) offers a fairly good approximation to the evolution of the labor share between 1960 and 1999 for most countries in our sample. Numerical results are reported in Table 3. As this table shows, inflation has a positive significant effect on the labor share in all countries, with the only exception of Austria and Denmark where inflation does not have a statistically significant impact.

Results for the rest of variables are less uniform. The output gap has a negative statistically significant coefficient in eight countries, and it is not a statistically significant variable in the remaining seven countries. The unemployment rate has a significant negative sign in ten countries. On the other hand, it is not significant in three countries, and shows a positive significant sign in two cases. Finally, the effect of the capital-labor ratio on the labor share differs across countries.

4 Concluding Comments

Our results show a robust positive relationship between the level of inflation and the labor share. This relationship comes up in the pooled estimations, as well as in almost all country-specific estimations. Furthermore, our results confirm the negative relationship between the labor share and contemporaneous output gap usually found in the literature, and point -in pooled estimations and in most individual cases- to a significant negative relationship between the labor share and unemployment.

It may be hypothesized, however, that the positive impact of inflation on the labor share could lessen in the very long run. Inasmuch as inflation lowers margins and increases fixed costs, changes in the long run average level of inflation may bring about long run changes in the structure of some industries. These changes may help recovering to some extent the initial markups, and reducing the effects of inflation on the labor share. Whatever is the case, since changes in industries structure may take a fairly long time, any reasonable rate of time preference

would likely bring about a significant present value of the distributive impact of inflation.

The implications of these results for the political economy of monetary policy seem to be considerable, since they provide empirical support for the distinct concern shown by trade unions and different political parties with regard to price stability. They suggest that the fight against inflation, though likely beneficial for aggregate output and growth, may involve more controversial issues -at least in the short and medium run- than usually considered. This is rather unfortunate since the role of economists as political advisers becomes more embarrassing when economic policies are likely to have distributive effects, than when it is possible to offer free lunch policies providing Pareto improvements. On the other hand, these results help explaining the negative effect that inflation seems to have on investment and growth.

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Table 1

Results for Different EstimationMethods

	LSDV	SUR	IV_1	IV_2
π_t	0.92 (14.49)	0.71 (18.74)	0.86 (3.28)	0.86 (4.20)
u_{t-1}	−0.03 (−6.61)	−0.03 (−14.81)	−0.03 (−2.03)	−0.03 (−1.85)
y_t	−0.39 (−4.34)	−0.29 (−6.83)	−0.79 (−3.70)	−0.87 (−3.90)
k_t	0.04 (4.77)	0.04 (9.21)	0.04 (1.21)	0.04 (1.20)
Adj. R^2	0.77	0.76	–	–

Note: Dependent variable is labor share. All variables are in logs. The t-statistics are reported in parentheses below the coefficient estimates. The data correspond to an unbalanced panel of 15 OECD countries over period 1960-99. Total number of observations are 514 for the OLS and SUR estimation.For the Instrumental Variable estimation (IV_1,IV_2) the number of observations are 478 and 488 respectively. In the instrumental variable estimation we use White's Heteroskedasticity consistent covariance matrix.

Table 2

Robustness of the Inflation Coefficient for different Especifications (SUR Estimation)

π_t	0.71 (18.74)	0.78 (20.79)	0.69 (18.86)	0.82 (21.61)	0.78 (20.72)
u_{t-1}	−0.03 (−14.81)	-	−0.02 (−9.29)	-	-
y_t	−0.29 (−6.83)	-	-	−0.21 (−4.56)	-
k_t	0.04 (9.21)	-	-	-	0.00 (0.39)
Adj. R^2	0.76	0.74	0.75	0.74	0.74

Note. Dependent variable is labor share. The first column of results correspond to the baseline (i.e., the SUR estimation in Table 1). The rest of columns correspond to the alternative elimination of the rest of the explanatory variables but the inflation rate, in the estimated equation. All variables are in logs. The t-statistics are reported in parentheses below the coefficient estimates. The data correspond to an unbalanced panel of 15 OECD countries over period 1960-98. Total number of observations are 522 for the first and the third columns, 514 for the second and 521 for the fourth.

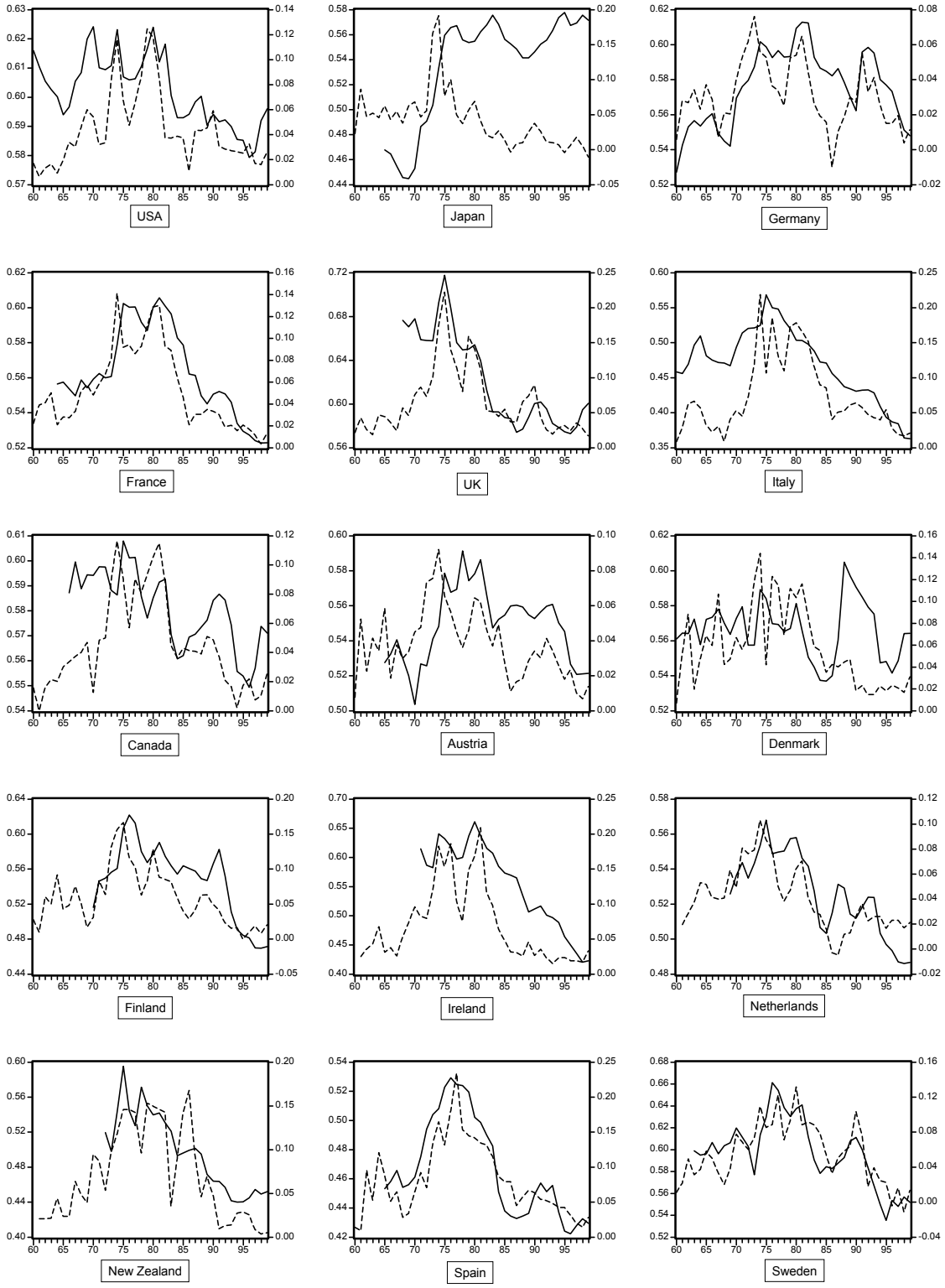
Table 3

Country-Specific Estimations
(SUR Estimation)

Country	*Constant*	π_t	u_{t-1}	y_t	k_t	*Adj.* R^2
USA	−0.18	0.40	−0.01	−0.24	−0.06	0.74
	(−4.12)	(9.12)	(−2.75)	(−4.37)	(−6.78)	
Japan	−1.23	0.47	−0.03	−0.94	0.15	0.92
	(−32.76)	(5.07)	(−1.30)	(−7.78)	(11.73)	
Germany	−0.87	1.01	−0.00	0.08	0.06	0.57
	(−12.74)	(5.13)	(−0.55)	(0.81)	(3.83)	
France	−0.38	0.93	0.03	−0.49	−0.06	0.75
	(−4.08)	(12.78)	(2.56)	(−4.63)	(−2.60)	
UK	−0.16	0.46	−0.06	−0.23	−0.06	0.95
	(−3.30)	(9.57)	(−8.76)	(−3.26)	(−4.49)	
Italy	−0.84	0.62	−0.27	−0.32	0.11	0.85
	(−9.25)	(4.74)	(−8.23)	(−1.16)	(3.51)	
Canada	−0.44	0.25	−0.04	−0.19	−0.01	0.63
	(−10.06)	(3.68)	(−4.48)	(−3.38)	(−0.81)	
Austria	−0.95	−0.00	−0.06	−0.55	0.09	0.32
	(−17.50)	(−0.03)	(−4.57)	(−3.73)	(6.26)	
Denmark	−0.65	0.03	−0.02	−0.34	0.02	0.14
	(−13.90)	(0.29)	(−3.08)	(−3.38)	(2.05)	
Finland	−1.02	0.47	−0.10	−0.18	0.11	0.59
	(−6.42)	(2.45)	(−6.34)	(−1.34)	(3.19)	
Ireland	0.22	1.35	0.23	0.14	−0.34	0.83
	(1.66)	(7.77)	(6.12)	(0.58)	(−8.35)	
Netherlands	−0.05	0.52	0.01	0.02	−0.13	0.46
	(−0.21)	(2.33)	(0.79)	(0.09)	(−2.45)	
New Zealand	−0.61	0.61	−0.04	−0.44	−0.02	0.84
	(−1.80)	(4.52)	(−4.99)	(−3.49)	(0.33)	
Spain	−1.07	0.70	−0.09	0.01	0.10	0.87
	(−17.37)	(9.21)	(−6.20)	(0.13)	(4.66)	
Sweden	−0.56	0.58	−0.06	−0.14	0.01	0.77
	(−9.95)	(4.65)	(−5.75)	(−1.09)	(1.01)	

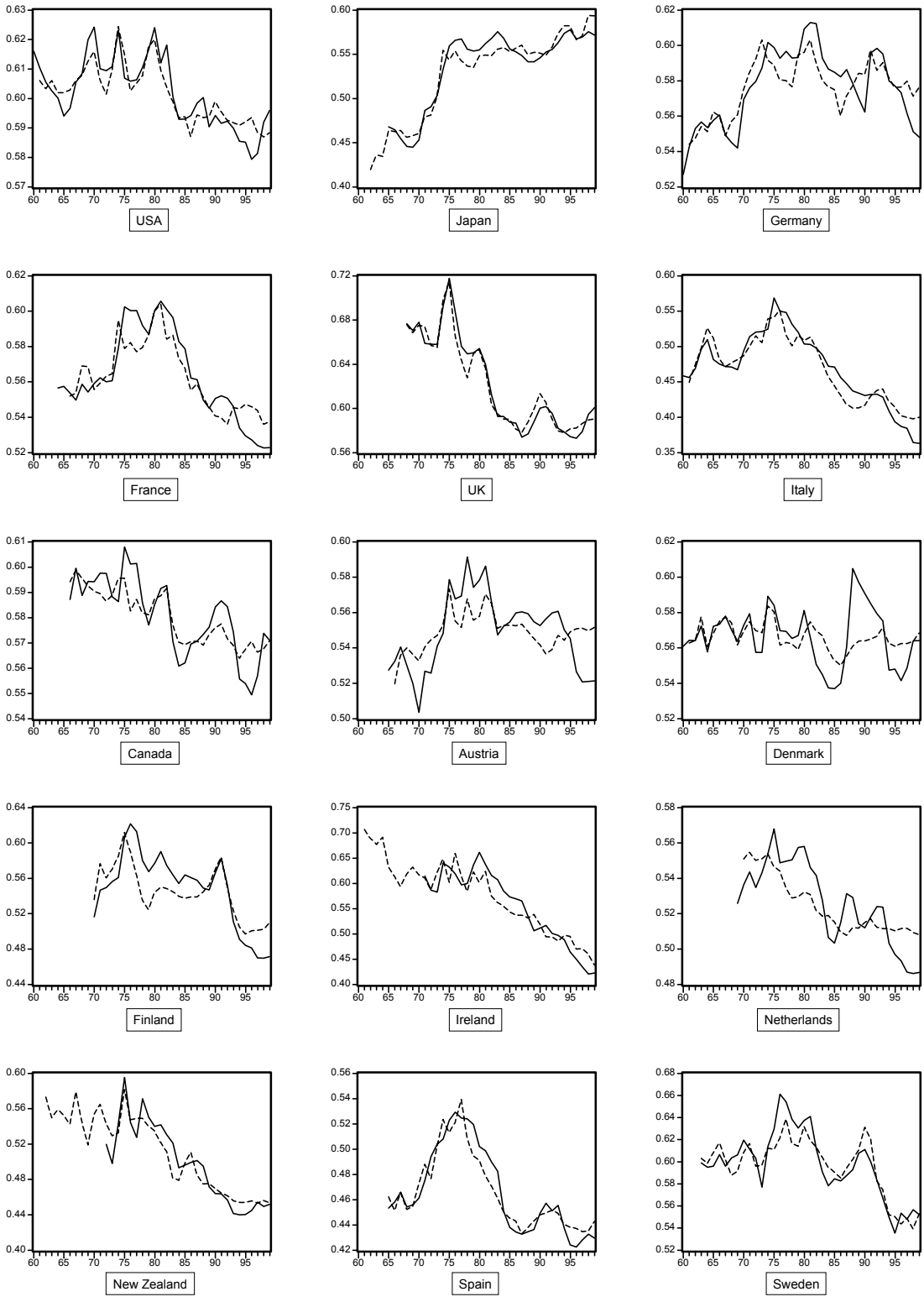
Note. Dependent variable is labor share. All variables are in logs. The t-statistics are reported in parentheses below the coefficient estimates. Errors for differents countries are assumed to be both contemporaneously correlated and cross-section heteroskedastic.

Figure 1. Inflation and labor share in OECD countries over the period 1960-99*



*Solid lines correspond to the labor share. Dashed lines correspond to the inflation rate.

Figure 2. Actual and fitted Labor share in OECD countries*



*Solid lines correspond to actual values. Dashed lines correspond to fits with equation (9)