Relative Price Variability and the Philips Curve: Evidence from Turkey

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Abstract

We argue that relative price changes are a key component of the Phillips curve relationship between inflation and output. Building on work by Ball and Mankiw, we propose including measures of the variances and skewness of relative price adjustment in an otherwise standard model of the Phillips curve. We examine the case of Turkey, where distribution of price changes is especially skewed and where the existence of a Phillips curve has been questioned. We have two main findings: (i) inclusion of measures of the distribution of relative price changes improves our understanding of the Phillips curve trade-off; (ii) there is no evidence of such a trade-off if these measures are not included.

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1) Introduction

Many studies have shown that consideration of the distribution of relative price adjustments can improve our understanding of the inflation rate. Early studies found a clear relationship between the level of inflation and the variance of relative prices (e.g. Vining and Elwertowski, 1976, Fischer, 1981, and Domberger, 1987). Following work by Ball and Mankiw (1994, 1995), more recent studies have also found a relationship between inflation and the skewness of relative price changes (e.g. Debelle and Lamont, 1997, Aucremanne et al., 2002 and Caraballo and Usabiaga, 2005). Although the relative size of the variance and skewness effects is controversial (e.g. Hall and Yates, 1988), the fact that the skewness effect appears quite strong for low inflation rates but much weaker when inflation is higher is consistent with the menu cost foundations of Ball and Mankiw's analysis.

In this paper we use these insights to improve our understanding of a key macroeconomic relationship, the Phillips Curve. We propose including measures of the distribution of relative price adjustment in an otherwise standard model of the Phillips curve. In doing so, we will combine two related but distinct literatures. The literature on the Phillips curve relates inflation to output or unemployment gaps. The literature on relative price variability relates inflation to the second and third moments of relative price changes. In this paper, we relate inflation to both factors.

We present empirical evidence for the case of Turkey. We do this for two reasons. First, the impact of the distribution of relative price changes on the Phillips curve may be more apparent in Turkey, where the distribution of relative price changes is markedly skewed. Second, there is some debate on whether the Phillips curve trade-off exists in Turkey (e.g. Kuştepeli, 2005; Önder, 2004 and Önder 2008). We hypothesise that this debate may reflect the difficulty in establishing a Phillips Curve if strong distributional effects from relative price changes are omitted from the model.

Beginning with a standard model of the hybrid Phillips curve similar to that derived by Gali and Gertler (1999), we first develop an empirical model in which inflation is determined by lagged values of inflation and current and lagged values of the output gap. We investigate the relationship between inflation, the output gap and the variance and skewness of relative price changes in Turkey, using monthly data for 1996:01 and 2007:05, for which we have information on prices of 75 sub-components of the consumer price index. We calculate standard measures of the standard deviation and skewness of changes in these disaggregated price indices, finding evidence of substantial skewness and variance and of marked changes in these distributional measures over time.

Our econometric approach is also a novelty in this literature. Since tests of the order of integration of our variables produced mixed results, we cannot be certain that all variables share the same order of integration. We therefore used the estimation procedure of Pesaran, Shin and Smith (1996, 2001) (hereafter, PSS). To do this, we estimated ARDL models in first differences, augmented by the lagged level values of our variables, with the differenced rate of inflation as the dependent variable. The bounds test procedure of PSS on the significance of these lagged terms was then used to assess whether the relationship is cointegrated. Estimates of any cointegrating relationships were then obtained by re-estimating this model expressed in terms of levels, with short-run dynamics being obtained by estimating the model in error-correction form.

Using this procedure, we find that the estimated relationship between inflation and the output gap is not cointegrated but that the relationship between inflation, the output gap and the variance and skewness of relative price changes is cointegrated. From this we conclude that there is a Phillips curve relationship in Turkey, but that omission of measures of the distribution of relative price changes can create the misleading impression that it does not.

The remainder of the paper is structured as follows. Section 2 provides an overview of past literature on relative price changes, inflation and the Turkish Phillips Curve and derives our empirical model. Section 3 describes our data and discusses the order of integration of our key variables and our estimation technique. Section 4 presents our econometric estimates and discusses their implications. Section 5 concludes.

2) Methodology

The literature on the relationship between inflation and the distribution of relative price changes typically estimates models of the form

$$(1) \qquad \pi_t = \beta_\pi(L)\pi_{t-1} + \beta_{sd}(L)sdrp_t + \beta_{sk}(L)skrp_t + \varepsilon_t$$

where π is the inflation rate, $sdrp$ is the standard deviation of relative price changes, $skrp$ is the skewness of relative price changes, ε is an iid error term, β_π, β_{sd} and β_{sk}, are polynomials of length n_π, n_{sd} and n_{sk} respectively in the lag operator L, where $\beta_\pi(L) = \beta_1^\pi + \beta_2^\pi L + ... + \beta_{n_\pi}^\pi L^{n_\pi - 1}$, $\beta_{sd}(L) = \beta_0^{sd} + \beta_1^{sd} L + ... + \beta_{n_{sd}-1}^{sd} L^{n_{sd}-1}$ and $\beta_{sk}(L) = \beta_0^{sk} + \beta_1^{sk} L + ... + \beta_{n_{sk}-1}^{sk} L^{n_{sk}-1}$.

Early studies (e.g. Vining and Elwertowski, 1976, Parks, 1978, Fischer, 1981, Domberger, 1987 and Hartman, 1991) examined the empirical relationships between inflation and relative price variability. Theoretical support for these relationships was provided Fischer (1981, 1982) and Cuckierman (1983). Following work by Ball and Mankiw (1994, 1995), who argued that, in the context of a menu cost model, an asymmetric pattern of relative price changes at the microeconomic level had implication for the behaviour of the aggregate inflation rate, the third moment of relative price changes was also considered (Balke and Wynne, 2000, argue that these effects can also arise in a model without price rigidities). This more recent literature has continued to find a strong association between inflation and the distribution relative price changes, although there is debate about the relative strength of the effect of the second and third moments. Some studies find that the effect of skewness is stronger (e.g. Ball and Mankiw, 1995, Debelle and Lamont, 1997, for the US; Aucremanne *et al.*, 2002, for Belgium; Caraballo and Usabiaga, 2005, for Spain), while De Abreu *et al.* (1995) for Australia; Bonnet *et al.* (1999) for France; Dopke and Pierdzioch (2003) for Germany and Assorson (2004) for Sweden, found the effects to be of roughly equal size. However some studies have found more ambiguous effects (see, for example, Hall and Yates (1998), for the UK; Ratfai (2004) for Hungary and Pou and Dabus (2005) for Spain and Argentina). More skeptical commentators include Holly (1997), who uses Japanese data to argue that causation runs from aggregate inflation to the distribution of relative price changes, and not vice-versa and Bryan and Cecchetti (1999), who argue that the relationships estimated in the literature reflect measurement error (but see, the rejoinder by Ball and Mankiw, 1999). It has also been suggested that a relationship based on menu-cost arguments will not be applicable in a context of a higher inflation rate where menu costs are less relevant.

Studies on Turkish data include Alper and Ucer (1998), who used a measure of relative price variability based on 21 subcomponents of the wholesale price index (WPI) for the 1985-97 period. The effect of relative price variability was not significant and there was no evidence that relative price variability has a Granger-causal relationship with the aggregate inflation rate. By contrast, Caglayan and Filiztekin (2001), using annual data from 1948 to 1997 found a strong relationship between relative price variability and the inflation rate, as did Kucuk and Tuger (2004) using monthly data for 1994-2002. To our best knowledge there appears no study which has examined the relationship between inflation and the third moment of relative price changes.

In this paper, we investigate whether the distribution of relative price changes affects the Phillips curve. This is not entirely novel, as some papers have included measures of unemployment or the output gap in equation similar to (1). However they are included as additional control variables and to check on the robustness of the relationship between inflation and the distribution of relative price changes (Dopke and Pierzdioch, 2001, include the unemployment rate in a model similar to (1), while Assarsson, 2004, includes unemployment relative to the natural rate of unemployment as one of eight control variables). To our knowledge, ours is the first paper systematically to investigate this issue.

We begin with the “hybrid” model of the Phillips curve, proposed by Gali and Gertler (1999), given by

(2) $$\pi_t = (1-\theta)\pi_{t-1} + \theta\delta E_t \pi_{t+1} + \gamma mc_t$$

where mc is the proportional deviation of marginal cost from it's steady-state value, δ is the discount rate and θ captures the relative weight on forward-looking price-setting. Gali and Gertler (1999) derive (2) using the Calvo (1983) model of nominal price adjustment but assuming that not all firms that are able to change price do so optimally, the other following a simple rule-of-thumb. The parameter θ reflects both the probability of being able to adjust price and the proportion of firms who reset prices optimally. Recent work has attempted to derive Phillips curves similar to (2) in the context of menu cost models (Gertler and Leahy, 2005) and information cost models (Mankiw and Reis, 2002), although models based around the Calvo model remain dominant (Dennis, 2007).

Since this paper uses time series techniques, it is convenient to express this model as

(3) $$\Delta\pi_t = -\frac{\theta(1-\delta)}{1-\theta\delta}\pi_{t-1} + \frac{\theta\delta}{1-\theta\delta}E_t\Delta\pi_{t+1} + \frac{\gamma}{1-\theta\delta}mc_t$$

We assume that expected future changes in the inflation rate can be expressed as a function of current and lagged inflation rates, $E_t\Delta\pi_{t+1} = \lambda_\pi(L)\Delta\pi_t$, where $\lambda_\pi(L) = \lambda_\pi^1 + \lambda_\pi^2 L + ... + \lambda_\pi^{n_\pi} L^{n_\pi - 1}$. We also assume that marginal cost can be expressed as a function of the output gap, $mc_t = \lambda_y(L)y_t$, where $\lambda_y(L) = \lambda_y^1 L + \lambda_y^2 L^2 + ... + \lambda_y^{n_y} L^{n_y}$. Substituting these into (3) yields

(4) $$\Delta\pi_t = -\bar{\lambda}_\pi \pi_{t-1} + \bar{\lambda}_{\Delta\pi}(L)\Delta\pi_{t-1} + \bar{\lambda}_y y_{t-1} + \bar{\lambda}_{\Delta y}(L)\Delta y_{t-1} + \varepsilon_t^s$$

where $\bar{\lambda}_\pi(L) = \dfrac{\theta(1-\delta)}{1-\theta\delta(1+\lambda_1^\pi)}$, $\bar{\lambda}_y(L) = \dfrac{\gamma}{1-\theta\delta(1+\lambda_1^\pi)}(\lambda_y^1 + \lambda_y^2 + ... + \lambda_y^{n_y})$,

$\bar{\lambda}_{\Delta\pi}(L) = \bar{\lambda}_{\Delta\pi}^1 + \bar{\lambda}_{\Delta\pi}^2 L + .. + \bar{\lambda}_{\Delta\pi}^{n_\pi} L^{n_\pi - 1} = \dfrac{\theta\delta}{1-\theta\delta(1+\lambda_1^\pi)}(\lambda_2^\pi + \lambda_3^\pi L + ... + \lambda_{n_\pi}^\pi L^{n_\pi - 1})$,

$\bar{\lambda}_{\Delta y}(L) = \bar{\lambda}_{\Delta y}^1 + \bar{\lambda}_{\Delta y}^2 L + .. + \bar{\lambda}_{\Delta y}^{n_y} L^{n_y - 1} = -\dfrac{\gamma\lambda_y^1}{1-\theta\delta(1+\lambda_1^\pi)}(\sum_{i=1}^{n_y}\lambda_i^y + \sum_{i=2}^{n_y}\lambda_i^y L + \sum_{i=3}^{n_y}\lambda_i^y L^2 + ...)$

and ε^s is an iid error term reflecting expectational errors. This model is the empirical counterpart of the hybrid Phillips curve in (2).

We next add measures of the second and third moments of relative price changes[1], giving the augmented Phillips curve

(5) $$\begin{aligned}\Delta\pi_t &= -\bar{\lambda}_\pi \pi_{t-1} + \bar{\lambda}_{\Delta\pi}(L)\Delta\pi_{t-1} + \bar{\lambda}_y y_{t-1} + \bar{\lambda}_{\Delta y}(L)\Delta y_{t-1} + \\ &\bar{\lambda}_{sd} sdrp_{t-1} + \bar{\lambda}_{\Delta sdrp\pi}(L)\Delta sdrp_{t-1} + \bar{\lambda}_{sk} skrp_{t-1} + \bar{\lambda}_{\Delta skrp}(L)\Delta skrp_{t-1} + \varepsilon_t^s\end{aligned}$$

where $\beta_{sd}(L) = \lambda_{sd}^1 + \lambda_{sd}^2 L + ... + \lambda_{sd}^{n_{sd}} L^{n_{sd}-1}$ and $\beta_{sk}(L) = \lambda_{sk}^1 + \lambda_{sk}^2 L + ... + \lambda_{sk}^{n_{sk}} L^{n_{sk}-1}$. Our empirical strategy will be to estimate the ARDL models in (4) and (5) and test whether the augmented model in (5) is superior. As with other models in the literature, there are no formal micro-foundations for (4). This is beyond the

[1] We did not include the cross product of $skrp$ and $sdrp$, as in Ball and Mankiw (1995), because of multicollinearity.

scope of this paper, but we would speculate that these will emerge once the literature has produced menu cost models that can generate Phillips curve models similar to (4). Drawing on the more heuristic microfoundations provided by the work of Ball and Mankiw (1994, 1995), we expect $\overline{\lambda}_{\pi} > 0$, $\overline{\lambda}_{y} > 0$ $\overline{\lambda}_{sd} > 0$ and $\overline{\lambda}_{sk} > 0$.

3) Data

We use monthly Turkish data for the period 1996:01 and 2007:05. The inflation rate is the proportional month-on-month change in the Index of Consumer Prices (HICP) (taken from the Eurostat database). The output gap is the proportional difference of de-seasonalised real GDP (made available by the Central Bank of the Republic of Turkey) from its' underlying Hodrick-Prescott (1992) trend.

Figure 1 depicts the inflation rate and output gap over the sample period. As can be seen from the figure Turkey has experienced high inflation accompanied by volatile growth until the end of 2002. In an attempt to end a long sequence of high inflation rates, an IMF-directed disinflation program, based on nominal exchange rate stability, was adopted in the beginning of the 2000. Eleven months later, this program was abandoned in the face of an economic crisis triggered by banking sector fragility and accumulating current account deficits, in favour of floating exchange rate regime (see, Alper, 2001, and Akyurek, 2006 for details). A rapid and depreciation of the Lira followed (the currency lost 51 percent of its value against major currencies), which led to a monthly inflation rate of 11.8 percent by April 2001 and an annual inflation rate of 75.1 percent in 2001. Following these traumas, the Central Bank of Turkey adopted a policy of monetary base targeting in early 2002, with an explicit focus on lowering and then stabilising the future inflation; this was in effect a regime of implicit inflation targeting but where the main policy instrument was the monetary base. This policy has proved successful. Inflation gradually decreased throughout 2002 and has remained largely low and stable since.

We use data on 75 sub-components of the price index[2]**.** The individual rate of inflation of each of these sub-components is calculated as

(6) $$\pi_{i,t} = p_{it} - p_{it-1}$$

where p_{it} is the natural logarithm of the price of sub-component *i* at time *t* and where the aggregate price is defined as $\pi_t = \sum_{i=1}^{N} w_i \pi_{i,t}$, where w_i is the weight on sub-component *i, where i*=1,…,75[3] . We use standard measures of the distribution of relative price changes. The second moment is defined as

[2] Some of the sub-components were not available for the whole sample period, therefore we used main components for these items and hence reduced the data to 75 subcomponents.
[3] The data related to 1996-2007 weights of the CPI was not fully available; therefore we used 1996 weights in this study.

(7) $$sdrp_t = \sqrt{\sum_{i=1}^{N} w_i \left(\pi_{i,t} - \pi_t\right)^2}$$

while the third moment is defined as

(8) $$skrp_t = \frac{\sum_{i=1}^{N} w_i \left(\pi_{i,t} - \pi_t\right)^3}{sdrp_t^3}$$

Figure 2 depicts $sdrp$ and $skrp$. Relative price changes are clearly highly volatile. Movements in the second moment are move with changes in the inflation rate. This closely relationship has been widely documented in previous studies (see, for example, Ball and Mankiw (1995), Debelle and Lamont (1997), Aucremanne *et al.* (2002), Caraballo and Usabiaga (2005), De Abreu *et al.* (1995), Bonnet *et al.* (1999), Dopke and Pierdzioch (2003) and Assorson (2004), Hall and Yates (1998), Ratfai (2004), Pou and Dabus (2005)). However we note that the reduced inflation rate in recent years has only partially been reflected in lower volatility**.** The skewness of relative price changes is most marked in periods of macroeconomic stress, when larger negative values are apparent. Overall, skewness has reduced in recent years.

4) Econometric Estimates

We begin by examining the stationarity properties of our data. As Table 1 shows, application of a variety of tests produces mixed results. We therefore use the bounds testing procedure proposed by Pesaran, Shin and Smith (1996, 2001) which allows us to test for the existence of a linear long run relationship with variable which may be of differing orders of integration.

To do this, we first estimate the ARDL models in (4) and (5) using ordinary least squares. We then test the restriction that all estimated coefficients of lagged variables equal zero by means of an F-test. In the case of (4), the null hypothesis of no cointegration corresponds to $H_0 : \overline{\lambda}_\pi = \overline{\lambda}_y = 0$. For (5) the null is $H_0 : \overline{\lambda}_\pi = \overline{\lambda}_y = \overline{\lambda}_{sd} = \overline{\lambda}_{sk} = 0$. This test has a non-standard asymptotic distribution, for which PSS provide two sets of critical values, corresponding to the cases where all variables are I(0) and where all variables are I(1). These upper and lower bounds constitute a range that includes all possible combinations of I(1), I(0) (or even fractionally integrated) variables. If the F-statistic lies above the upper critical bound, the null of no cointegration is rejected, while the test is inconclusive if the F-statistic lies between the upper and lower bounds. Any long run relationship that is detected can then be estimated using an ARDL model similar to (4) and (5) above but which includes lags of the levels rather than the first differences of the variables of interest. Short-run dynamics can then be obtained by estimating an error correction version of this model, where the estimated long-run relationship forms the error-correction term.

We estimated the conditional ARDL models using up to 13 lags, (although we only included one lag of $sdrp_t$; further lags were not significant and were omitted to prevent over-parameterisation). We also included a dummy variable for April 2001, which was interacted with the output gap to correct for a sharp and anomalous drop in output in that month (at the height of the crisis of early-mid 2001). For each model, we calculated tests of serial correlation, since, as PSS point out, the validity of these tests for cointegration requires serially uncorrelated residuals.

Cointegration tests for the model in (4) are presented in Table 2. As column (v) of that table shows, the test statistic exceeds the upper critical value in the case where 3 lags are used. However, as column (iv) shows, that model suffers from serial correlation. The test statistic is in the inconclusive zone when 1 or 2 lags are used, but these models also fail the test for serial correlation. In all other cases, the test statistic for cointegration is less than the lower critical value. Therefore the null hypothesis of no cointegration in estimates of (4) is never rejected. In other words the Phillips curve relation is not valid for Turkey, casting doubt on this fundamental macroeconomic relationship. There is some debate on the existence of the Turkish Phillips Curve in the literature. While Kustepeli (2005) finds no evidence of a Phillips curve in Turkey, Önder (2004) founds a linear relationship by using output gap instead of unemployment gap. On the other hand, Önder (2008) investigates instability of the Phillips curve and she finds weaker support for the curve by taking nonlinearities into account

Tests for the model in (5) are presented in Table 3. The results in this case are very different as there is strong evidence that the augmented Phillips curve model in (5) is cointegrated. The null hypothesis of no cointegration is rejected in every model that does not from serial correlation. Inclusion of the higher moments of the distribution of relative price changes has allowed the Phillips curve relationship to be established.

Having established that (5) is cointegrated, we estimated a levels version of (5), as discussed above[4], to extract estimates of this relationship. They are

$$\pi_t = \underset{(0.007)}{-0.02} + \underset{(0.079)}{0.228} y_t + \underset{(0.149)}{0.822} sdrp_t + \underset{(0.037)}{0.174} skrp_t \tag{8}$$

where standard errors are in parentheses. All estimated coefficients are significantly different from zero and have expected signs. The coefficients above do not represent elasticities and standard deviation and skewness differ in terms of magnitude (See Figure 1 and 2). Therefore we have calculated average elasticity of inflation with respect to skewness and

[4] We included a full lag structure for $skrp$, as suggested by PSS. The specification of our ARDL was determined by the AIC criteria, by which measure an ARDL(11,3,4,11) model performed best.

standard deviation and found as 3.45 and 1.30 respectively[5]. That means the effect of third moment of relative price variability is higher than that of standard deviation. This result is also consistent with Ball and Mankiw's result.

Finally, Table 4 presents estimates of the ARDL model expressed as an error-correction model and using the estimated cointegrating relationship as the error-correction term. The model passes diagnostic checks for normality, autocorrelation, misspecification and heteroscedasticity. Furthermore, Cumulative Sum of Residuals (CUSUM) and Cumulative Sum of Squared Residuals (CUSUMSQ) tests (these are not reported, but are available upon request) find no evidence of instability in the estimated coefficients. The error correction coefficient is large (-0.398) and highly significant. We estimate that 40% of the deviation from the long-run equilibrium level of inflation is corrected within a month. Although the dynamic structure is quite complex, it is apparent that almost all lags of skewness are very significant and the skewness of the underlying distribution of prices is a more persistent determiner of movements in variables at the macroeconomic level than is relative price variability. This suggests that the relative importance of skewness, first established by Ball and Mankiw (1995) in the context of (1), also applies in the case of the Phillips curve.

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[5] Elasticities are calculated by using the following formula $\varepsilon_{y,x} = \frac{\Delta y}{\Delta x} \cdot \frac{\overline{x}}{\overline{y}}$.

5) Conclusions

This paper has argued that relative price changes are a key component of the Phillips curve relationship between inflation and output. We have combined the literature on the relationship between inflation and the distribution of relative price changes with the literature on the Phillips curve by including the variance and skewness of relative price adjustment in an otherwise standard model of the Phillips curve. We examine the case of Turkey, where distribution of price changes is especially skewed and where the existence of a Phillips curve has been questioned.

We find that measures of the distribution of relative price changes do indeed improve our understanding of the Phillips curve trade-off. Using monthly data from 1996-2007, we find no evidence of a trade-off between inflation and output in a conventional model of the Phillips curve. By contrast, a well-determined trade-off is obtained when the variance and skewness of relative price changes is included in the model.

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Figure 1 – Consumer Price Inflation and Output Gap in Turkey: 1996:2-2007:5

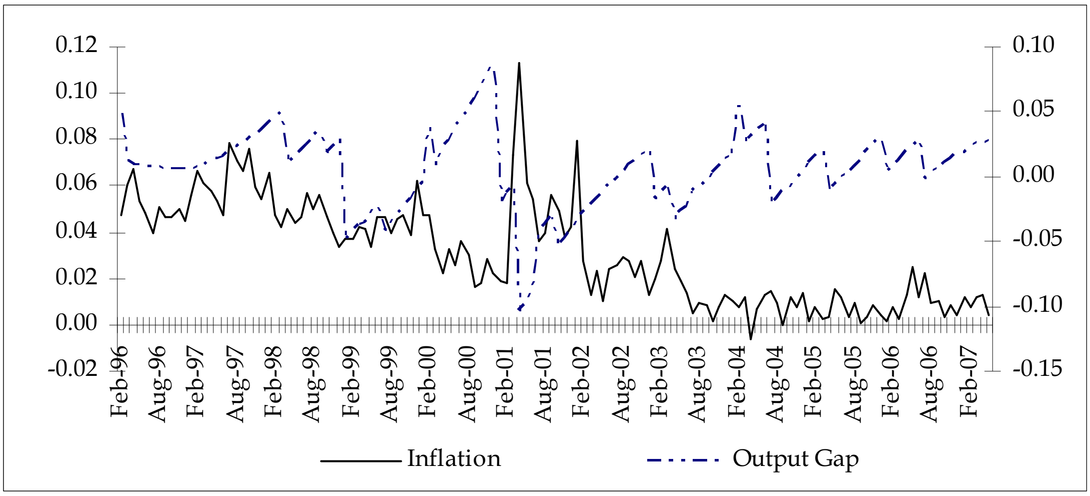

Figure 2- Standard Deviation of Relative Price Changes and Inflation in Turkey: 1996:2-2007:5

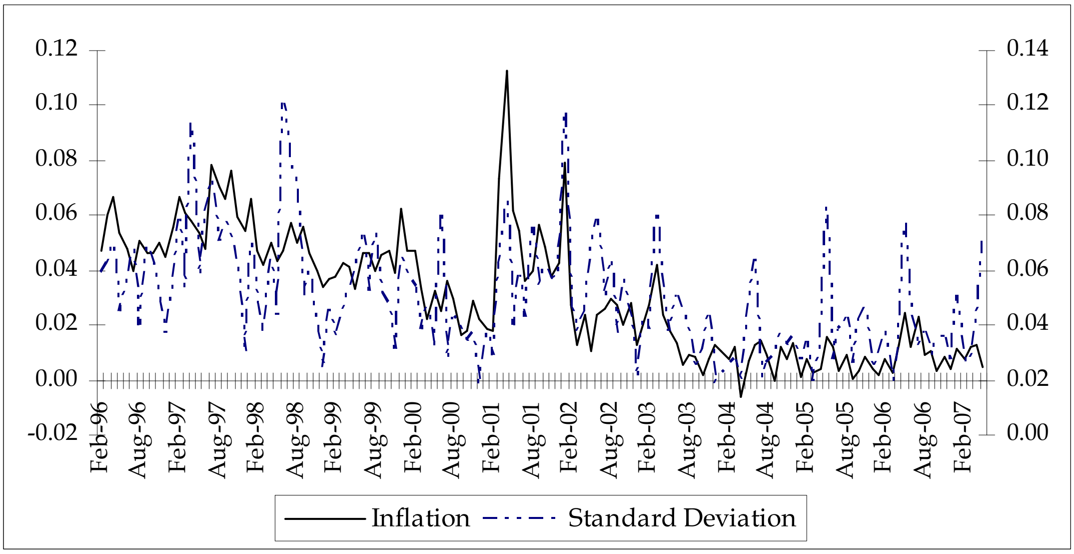

Figure 3- Skewness of Relative Price Changes and Inflation in Turkey: 1996:2-2007:5

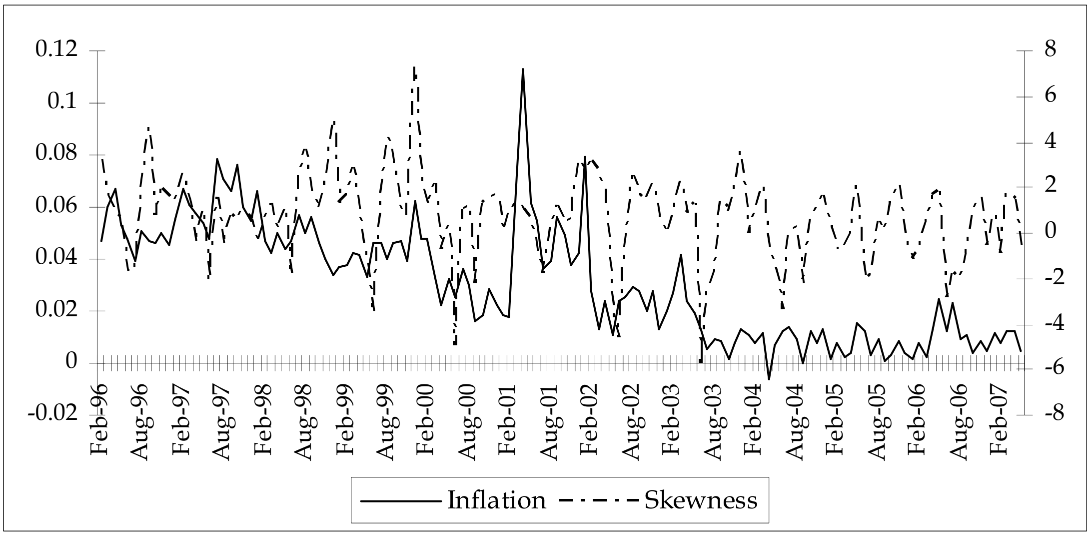

Table 1: Unit Root Tests

	ADF	PP	KPSS	DFGLS	$NGP_{(MZ\alpha)}$
π	-6.175***	-6.105***	1.252***	-2.356	-14.49*
Δπ	-	-	0.220	-13.648***	-
y	-3.544***	-3.986***	0.115*	-3.389	-20.336
Δy	-	-	-	-0.822*	-0.525*
sdrp	-1.38	-9.262***	0.065	-1.00*	-2.579*
Δsdrp	-5.47***	-	-	-	-
skrp	-2.963	-8.184***	0.561	-0.100 *	0.235*
Δskrp	-9.728***	-	-	-	-

Note: *, ** and *** indicate significant at 10, 5 and 1% respectively. The lag length for ADF test is chosen based on the AIC criterion. Contrary to other unit root tests null hypothesis of KPSS test is stationary. Bandwiths in the PP and KPSS unit root tests are determined by the Newey-West statistic using the Barlett-Kernel. The lag length of the DF-GLS and Ng-Perron tests are selected by the Modified Akaike Information Criterion (MAIC).

Table 2: Bounded F-tests for Phillips Curve for model (4)

Lag	AIC	SBC	$\chi^2_{sc}(12)$	F-statistics
1	397.623	388.908	39.4574(.000)	4.764 (i)
2	402.094	391.952	25.9915(.011)	4.615 (i)
3	398.269	385.262	29.6960(.003)	4.900 (r)
4	398.843	382.987	25.4362(.013)	3.278
5	393.697	375.008	29.2646(.004)	2.647
6	388.885	367.378	25.8388(.011)	2.689
7	384.689	360.381	29.4465(.003)	2.519
8	382.468	355.373	25.9342(.011)	1.811
9	378.866	349.002	27.7414(.006)	2.301
10	374.798	342.181	30.6840(.002)	1.323
11	376.043	340.689	27.2604(.007)	0.446
12	373.018	334.944	20.9068(.052)	0.480
13	371.121	330.344	21.1679(.048)	0.669

Note: Asymptotic critical values for bounded F-test are 3.79 and 4.85 for I(0) and I(1) respectively 5% significance level. $\chi^2_{sc}(12)$ is LM test statistics for testing no serial correlation, p-values are in parenthesis. In column (v), (i) indicates a test statistic in the inconclusive range, while (r) indicates rejection of the null

Table 3: Bounded F-Tests For Phillips for model in (5)

Lag	AIC	SBC	$\chi^2_{sc}(12)$	F-statistics
1	388.558	371.685	26.2965(.010)	2.895 (i)
2	392.511	370.013	19.5594(.076)	3.568 (i)
3	391.396	367.493	17.1983(.142)	5.890 (r)
4	390.665	362.543	21.3265(.046)	4.9011 (r)
5	390.870	358.530	20.9821(.051)	5.738 (r)
6	389.252	352.6932	22.1253(.036)	4.250 (r)
7	387.110	346.333	23.3544(.025)	4.369 (r)
8	385.870	340.875	23.0645(.027)	4.745 (r)
9	389.814	340.601	20.9203(.052)	6.333 (r)
10	390.936	337.505	16.094(.207)	5.792 (r)
11	389.178	331.528	17.9594(.117)	4.396 (r)
12	388.812	326.944	14.0916(.295)	4.724 (r)
13	390.785	324.699	20.3149(.061)	4.922 (r)

Note: Asymptotic critical values for bounded F-test are 2.86 and 4.01 for I(0) and I(1) respectively at 5% significance level. $\chi^2_{sc}(12)$ is LM test statistics for testing no serial correlation, p-values are in parenthesis. In column (v), (i) indicates a test statistic in the inconclusive range, while (r) indicates rejection of the null hypothesis.

Table 4: Error Correction Form of the ARDL(11,2,11,12) Phillips Curve Model

Regressor	Coefficient	Standard Error	p-value
Δπ(-1)	-0.212	0.101	0.039
Δπ(-2)	-0.165	0.099	0.099
Δπ(-3)	-0.023	0.093	0.807
Δπ(-4)	0.031	0.088	0.723
Δπ (-5)	0.175	0.086	0.044
Δπ (-6)	0.213	0.086	0.015
Δπ (-7)	0.181	0.080	0.027
Δπ(-8)	0.144	0.079	0.071
Δπ(-9)	0.312	0.072	0.000
Δπ(-10)	0.173	0.070	0.015
Δy	0.005	0.045	0.916
Δy(-1)	-0.120	0.045	0.009
Δy(-2)	-0.183	0.043	0.000
Δsdrp	0.315	0.043	0.000
Δsdrp(-1)	0.072	0.074	0.335
Δsdrp(-2)	0.072	0.064	0.263
Δsdrp(-3)	0.105	0.048	0.032
Δskrp	0.002	0.000	0.000
Δskrp (-1)	-0.005	0.001	0.002
Δskrp(- 2)	-0.004	0.001	0.003
Δskrp (-3)	-0.004	0.001	0.001
Δskrp (-4)	-0.003	0.001	0.003
Δskrp (-5)	-0.003	0.001	0.006
Δskrp(-6)	-0.003	0.001	0.000
Δskrp(-7)	-0.003	0.001	0.000
Δskrp (-8)	-0.002	0.001	0.002
Δskrp (-9)	-0.003	0.001	0.000
Δskrp(-10)	-0.001	0.000	0.012
Constant	-0.009	0.003	0.004
Dummy	-0.633	0.098	0.000
Ecm(-1)	-0.398	0.082	0.000
R-Bar-Squared	0.765		
F-stat. F(36, 88)	13.356(.000)		
$\chi^2_{SC}(12)$	.10446(.747)	$\chi^2_{H}(12)$	8.8177[.718]
$\chi^2_{FF}(1)$	1.9868(.159)	$\chi^2_{N}(12)$	

Notes: $\chi^2_{SC}(12)$, $\chi^2_{H}(12)$, $\chi^2_{FF}(1)$ and $\chi^2_{N}(12)$ denote chi-squared statistics for residuals, to test the null hypothesis of no serial correlation, no functional form misspecification, normality and homoscedasticity respectively. p values are in parenthesis.