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Scientific review

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A Cost-based Empirical Model of the Aggregate Price Determination for the Turkish Economy: A Multivariate Cointegration Approach

Summary: This paper tries to examine the long run relationships between the aggregate consumer prices and some cost-based components for the Turkish economy. Based on a simple economic model of the macro-scaled price formation, multivariate cointegration techniques have been applied to test whether the real data support the a priori model construction. The results reveal that all of the factors, related to the price determination, have a positive impact on the consumer prices as expected. We find that the most significant component contributing to the price setting is the nominal exchange rate depreciation. We also cannot reject the linear homogeneity of the sum of all the price data as to the domestic inflation. The paper concludes that the Turkish consumer prices have in fact a strong cost-push component that contributes to the aggregate pricing.

Key words: Aggregate consumer prices, Cointegration, Turkish economy.

JEL: C32, C51, E31.

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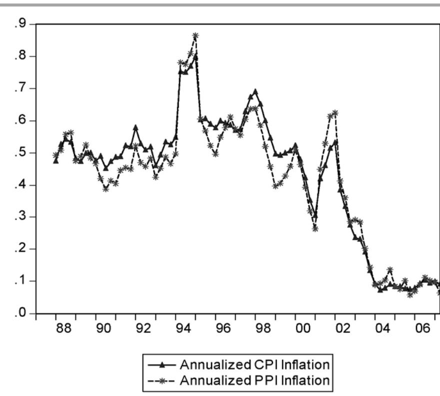
A chronic inflationary framework is one of the main properties that has identified the course of the Turkish business cycles for the last 30-years period, and constitutes an important benchmark for economic agents in constructing their expectations. The data from the post-1980 period indicate that inflation rate took annual values within the range of 30% - 50% for the years 1981 - 1987. Following this sub-period, the economy witnessed a jump in annual inflation, and inflation rates began to fluctuate between 60% - 80%. Conditions of the 1994 economic crisis led to a one-time upward jump in annual inflation rates and inflation lay between 80% - 100% interval for the 1995 - 1998 period. For the post-1998, annual inflation followed a downward trend, however, it remained above the 55% - 60% minimum threshold levels of the previous periods till the year 2000. As of the year 2000, the Turkish economy embarked on an anti-inflationary stabilization program based on a crawling peg/band regime to fight domestic inflation, and policy makers aimed at mainly forming the expectations of economic agents in line with the policy issues consistent with nominal exchange anchor.[1] Although seemed to be successful in bringing inflation down instantly to the 35% annual level for the first 10 months of realization, the subsequent two economic crisis periods ended the program. Following the collapse of the nominal exchange anchor based disinflation stabilization program, a massive economic crisis took place in 2001, that led to a great slump in real income by about -9.50%, and in turn this period coincided with an upsurge of annual inflation within the range of 60% - 65%.

For the post-2002 period, policy makers decided to establish an inflation targeting framework that was applied implicitly for the pre-2006 period under the acceptance of the indepencence of the monetary authority in implementation of monetary stabilization policies. Hakan A. Kara (2006) describes the challenges faced during the implementation of implicit inflation targeting in Turkey in a highlighting way and evaluates the transition process to the full-fledged inflation targeting. The policy has turned out to be rather explicit targeting for the post-2006 period through the announcements of the annual targets determined in a co-ordinated way with the central government. In this period, annual inflation steadily dropped till the 8% - 10% threshold values and has been subject to an inertia to drop further. Thus the post-1980 experience of the Turkish economy indicates that inflation tends to mainly be characterized with the realizations of self-peculiar characteristics as to the sub-periods, rather than with a stable long-run path. On this point, see also Ahmet Ertuğrul and Faruk Selçuk (2002) for a brief outline of the Turkish economy considering the whole 1980s and 1990s.

We have summarized the development of the Turkish inflation in Figure 1 below. In the figure, we present the annualized consumer prices (CPI) and producer prices (PPI) inflation with the base 2000: 100. We can easily observe the volatility Turkish inflation indicates from the late 1980s till the mid-2007.



Source: Organization for Economic Co-operation and Development (OECD) Statistics Portal[2].

Figure 1 The Development of Inflation in Turkish Economy

[1] For details of the Letter of Intent that Turkey declared her targets, see http://www.imf.org/external/NP/LOI/1999/120999.htm.

[2] **Organization for Economic Cooperation and Development.** 2010. OECD.Stat Extracts. http://stats.oecd.org/index.aspx (accessed February 2, 2010).

There exists a large literature constructed upon the reasons of this issue of interest for the Turkish economy. In this respect, G. C. Lim and Laura Papi (1997) observe that monetary factors play a central role in the inflationary process and that public sector deficits significantly contribute to the inflation. They conclude that the inertial factors are quantitatively important for the Turkish inflation. Pierre R. Agénor and Alexander W. Hoffmaister (1997) find that the primary role in the movement of inflation for the Turkish economy can be attributed to the innovations in inflation itself and the innovations in exchange rate depreciation. Emre C. Alper and Murat Üçer (1998), Bilin Neyaptı (1998), Cem Akyürek (1999), Christopher F. Baum, John Barkoulas, and Mustafa Çağlayan (1999), and Kıvılcım Metin-Özcan, Hakan Berument, and Neyaptı (2004) emphasize the importance of the strong inertial nature of the domestic inflationary framework, and generally attribute the nominal dimension of prices to the exchange rate depreciations and the policy framework following the real exchange rate rules applied in 1980s and 1990s. Ümit Cizre-Sakallıoğlu and Erinç Yeldan (1999) and Metin-Özcan, Ebru Voyvoda, and Yeldan (2001), using a business cycle framework, give supportive estimation results to such inferences for the Turkish consumer prices. Haluk Erlat (2002) also suggests that since inflation rates have a stationary characteristic with a significant long-memory component, the stabilization programs in fighting inflation must take account of high resistence in inflation rates. Erdal Özmen (1998) and Ayça Tekin Koru and Özmen (2003) find that in the long-run inflation appears to determine the currency growth and that inflation does not seem to be the result of an active monetary policy aiming to maximize seigniorage revenues. Likewise, Vuslat Us (2004) attributes the relatively high and inertial nature of the Turkish inflation to the increases in public sector prices and the depreciation of domestic currency, and indicates that high prices have not been as a result of expansionary monetary policy, leading to the conclusion that the inertial nature of the Turkish inflation is not a monetary phenomenon. Sel Dibooglu and Aykut Kibritcioglu (2004) emphasize that dis-inflation programs applied in the Turkish economy must have credible commitment mechanisms that restrain discretionary aggregate demand policies. Cem Mehmet Baydur and Bora Süslü (2004) estimate that the Central Bank of the Republic of Turkey (CBRT) assisted in the rise of inflation by implementing tight monetary policy from 1987 to 1997 and that it contributed to the fall of inflation by following relatively loose monetary policy after 1997. They also state that the CBRT does not have monopolistic power in controlling the inflation rates. Thus, we can infer here that the papers on the Turkish inflation tend to mainly emphasize the importance of the cost-based explanation, e.g. due to the exchange rate developments, and the inertial nature of the inflation.

In this paper, our contribution to the existing literature is to empirically examine the appropriateness of a cost-push model of the aggregate price-setting in the economy. To this end, a simple economic model has been developed and then tested in the light of some contemporaneous time series estimation techniques. For this purpose, the next section is devoted to the model construction. The second section describes the preliminary data issues and the third section tries to briefly highlight the methodological issues used in the model estimation process. The fourth section applies the multivariate cointegration techniques to test the data consistency of the theoretical model. Finally, the last section summarizes the results to conclude the paper.

1. A Simple Cost-Push Model for Aggregate Price Setting

In our paper, we tend to follow such papers as Gordon de Brouwer and Neil R. Ericsson (1995), Toshitaka Sekine (2001) and İlker Domaç (2004) to construct a cost based model in explaining the long-run course of the consumer prices in the Turkish economy. We assume that in a long-run perspective, the aggregate consumer prices level tend to be affected by some cost-factors, which are assumed to mainly be comprised of unit labor costs (*ULC*) as an index of the nominal costs of labor per unit of output, the nominal exchange rate (*E*) developments and the foreign prices (P^{for}) which are both assumed to reflect the amount of imported costs for the domestic economy. We have also included the domestic producer prices (P^{ws}) as an explaining factor of the consumer prices, since they are able to represent the course of the prices of inputs, such as intermediate goods and energy, determined in the earlier stages of the production of goods. In this way, we tried to incorporate them into the formation process of the consumer prices. For any given period *t*, we can write down such a pricing rule in a functional form as follows:

$$P_t = \mu_t (ULC_t)^{\gamma} . (E_t)^{\delta} . (P_t^{for})^{\eta} . (P_t^{ws})^{\phi} \tag{1}$$

In Eq. 1, the elasticities of the consumer prices with respect to unit labor costs, nominal exchange rate, foreign prices and the domestic producer prices are γ, δ, η and ϕ, respectively. These elasticities are hypothesized to be greater than or equal zero. If we use a log-linear form of Eq. 1, we can express it as follows, where the logarithms of the variables are denoted by lower case letters:

$$p_t = \ln(\mu_t) + \gamma . ulc_t + \delta . e_t + \eta . p_t^{for} + \phi . p_t^{ws} \tag{2}$$

For Eq. 2, we are simply able to test the linear homogeneity of the model as to the prices by applying to sum-of-coefficients restriction that amounts to a unit value:

$$\gamma + \delta + \eta + \phi = 1 \tag{3}$$

Following De Brouwer and Ericsson (1995), under the hypothesis of unit homogeneity in all prices, linear homogeneity allows us to re-write Eq. 2 as follows:

$$0 = \ln(\mu_t) + \gamma .(ulc_t - p_t) + \delta .(e_t - p_t) + \eta .(p_t^{for} - p_t) + \phi .(p_t^{ws} - p_t) \tag{4}$$

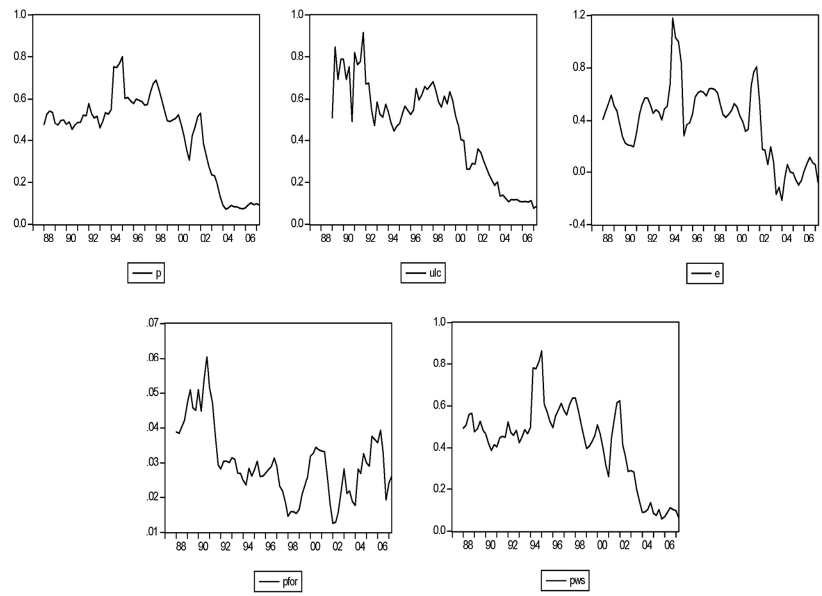
This formulation, if it can also be supported by the actual data, enables researchers to link the real prices of the various markets in the economy such as labor, foreign goods and input markets.[3] In our paper, we try to empirically test these relationships within a long-term perspective by applying to some contemporaneous time series estimation techniques.

2. Preliminary Data Issues

We now describe the data used in the paper and try to briefly highlight the method to test the empirical validity of the pricing model constructed in the former section. The sample considers the time period 1988Q1 - 2007Q2 with quarterly frequency data. All the variables are in their natural logarithms and have been converted to annual growth rates such that for any variable x_t observed at time t, $\Delta_4 x_t = (1-L^4)x_t$ where Δ is the difference operator defined as (1-L) and the lag operator L shifts x_t one period into the past. The domestic consumer price inflation variable (p_t) is derived from the 2000: 100 based consumer price index including all items in the price basket.[4] The annual growth of the unit labor costs (ulc_t) are represented by the 1997: 100 based index of wages per production hour worked in the manufacturing industry. For the annual nominal exchange rate depreciation data (e_t), the depreciation rate of the Turkish lira per US$ is considered. The annual change in foreign prices (p_t^{for}) data are from the 2000: 100 based consumer price index for the US economy. Finally the annual change in domestic producer prices (p_t^{ws}) reflect the 2000: 100 based producer price index data. The domestic and foreign price data are obtained from the electronic statistic portal of OECD, while the relevant wage and nominal exchange rate data have been taken from the electronic data delivery system of the Central Bank of the Republic of Turkey (CBRT). The time series graphs are reported below:

[3] On this point, see also Katarina Juselius (1992) and Metin (1995) that examine the effects of the price developments in various markets on the course of the aggregate price level.

[4] The consumer prices used for Turkey are mainly based on the 1994 consumer expenditure survey which has been subject to 5-year frequency updating. For combining prices to obtain lowest level indices as elementary aggregates, the average price of a sample of observations in the current period is compared to the average price of the sample period in the base period, and then these elementary aggregates are combined using some kind of index number formula and weights based on expenditure. In the case of the Turkish data, a standard Laspeyres type formulation is used to obtain such higher level aggregation price data. For further detailed methodological information upon data weighting and index calculation of the price indices, see:
Organization for Economic Cooperation and Development. 2010. OECD database. http://www.oecd.org/dataoecd/60/61/1947731.pdf (accessed February 2, 2010).



Source: Authors' own calculations.

Figure 2 Time Series Graphs

The spurious regression problem analyzed by Clive W. J. Granger and Paul Newbold (1974) indicates that using non-stationary time series steadily diverging from long-run mean leads to unreliable correlations within the regression analysis leading to unbounded variance process. However, for the mean, variance, and covariance of a time series to be constant over time, conditional probability distributions of the series must be invariant with respect to the time. David A. Dickey and Wayne A. Fuller (1981) provide one of the commonly used test methods known as the augmented Dickey-Fuller (ADF) test of detecting whether the time series data are of stationary form. This can be formulated for any x_t variable as follows:

$$\Delta x_t = \alpha + \beta t + (\rho - 1) y_{t-1} + \sum_{i=1}^{k} \eta_i \Delta x_{t-i} + \varepsilon_t \tag{5}$$

of which the null hypothesis is the presence of a unit root (ρ=1) against the alternative (trend) stationary hypothesis. For x_t to be stationary, (ρ-1) should be negative and statistically different from zero. The estimated ADF statistics are compared with the simulated James G. MacKinnon (1996) critical values. For the case of stationarity, we expect that these statistics must be larger than the critical values in absolute value and have a minus sign.

However, conventional unit root tests tend to be strongly criticized in the contemporaneous economics literature when they have been subject to structural breaks which yield biased estimations. These tests assume that variables can be characterized as a random walk process which requires differencing to achieve a stationary time series. Thus, we additionally follow the widely used Eric Zivot and D. W. K. Andrews (1992) (henceforth ZA) method allowing the data to indicate breakpoints endogenously rather than imposing a breakpoint from outside the system. Briefly to say, the ZA test chooses the breakpoint as the minimum t-value on the autoregressive x_t variable, which occurs at time 1 < TB < T leading to $\lambda = TB / T$, $\lambda \in \left|0.15, 0.85\right|$, by following the augmented regressions:

Model A: $$x_t = \mu + \beta t + \theta DU_t(\lambda) + \alpha x_{t-1} + \sum_{i=1}^{k} c_i \Delta x_{t-i} + \varepsilon_t \quad (6)$$

Model B: $$x_t = \mu + \beta t + \gamma DT_t^*(\lambda) + \alpha x_{t-1} + \sum_{i=1}^{k} c_i \Delta x_{t-i} + \varepsilon_t \quad (7)$$

Model C: $$x_t = \mu + \beta t + \theta DU_t(\lambda) + \gamma DT^*(\lambda) + \alpha x_{t-1} + \sum_{i=1}^{k} c_i \Delta x_{t-i} + \varepsilon_t \quad (8)$$

Above, DU_t and DT_t are sustained dummy variables capturing a mean shift and a trend shift occuring at the break date respectively. Δ is the difference operator, k is the number of lags determined for each possible breakpoint by one of the information criteria and ε_t is assumed to be an identically and independently distributed (i.i.d.) error term. The ZA method runs a regression for every possible break date sequentially and the time of structural changes is detected based on the most significant t-ratio for α. To test the unit root hypothesis, the smallest t-values are compared with a set of asymptotic critical values estimated by ZA. All of the unit root test results that led us to infer how integrate the variables are given in Table 1 and Table 2. Note that if x_t is found an I(k) process then $\Delta^k x_t$ is I(0).

The unit root test results from the ADF equation indicate that the null hypothesis cannot be rejected for all the variables in their levels, and differencing provides stationarity. Therefore we infer that all of the variables have an I(1) characteristic due to the ADF test results. When we consider the ZA unit root test results in Table 2 allowing one endogenous break in the time series used, no change occurs in the non-stationary characteristics of the variables.

Table 1. ADF Unit Root Tests

Variables	*in levels* τ_c^{ADF}	τ_t^{ADF}	*in first differences* τ_c^{ADF}	τ_t^{ADF}	*Inference*
p_t	0.24 (4)	-1.47 (4)	-6.58 (3)*	-6.93 (3)*	I(1)
ulc_t	-0.30 (4)	-1.74 (4)	-5.91 (3)*	-5.89 (3)*	I(1)
e_t	-1.24 (5)	-2.02 (5)	-4.39 (4)*	-8.11(3)*	I(1)
p_t^{for}	-2.60 (8)	-1.99 (8)	-4.17 (7)*	-4.51 (7)*	I(1)
p_t^{ws}	-0.10 (4)	-1.40 (4)	-7.23 (3)*	-7.41 (3)*	I(1)

Notes: τ_c and τ_t are the test statistics for the ADF tests with allowance for only constant and constant&trend terms in the unit root tests, respectively. 5% critical values are $\tau_{c,0.05}$=-2.90

and $\tau_{t,0.05}$=-3.47. * denotes the rejection of the unit root null hypothesis at the 5% level. The numbers in parentheses are the lags used for the ADF test, which are augmented up to a maximum of 10 lags. The choice of optimum lag for the ADF test was decided on the basis of minimizing the Schwarz information criterion.

Source: Authors' own calculations.

Table 2. ADF Unit Root Tests

	Intercept			*Trend*			*Both*		
	k	*min t*	*TB*	*k*	*min t*	*TB*	*k*	*mint*	*TB*
p_t	0	-4.041	02Q2	0	-3.210	94Q4	0	-4.217	94Q2
ulc_t	1	-3.473	95Q2	1	-2.715	98Q1	1	-3.462	95Q2
e_t	1	-4.016	02Q1	1	-4.407	94Q3	1	-4.873	94Q1
$p^{for}{}_t$	0	-4.017	91Q3	0	-3.311	94Q1	0	-4.195	91Q1
$p^{ws}{}_t$	0	-3.987	02Q1	0	-3.371	95Q1	0	-4.441	94Q2

Notes: Estimations with 0.15 trimmed. *min t* is the minimum *t*-statistic calculated. 5% critical values - intercept: -4.80; trend: -4.42; both: -5.08. min-*t* is the Schwarz Bayesian information criterion-minimizing value.

Source: Authors' own calculations.

3. Multivariate Co-integration Methodology

We now try to test for a long-run stationary relationship within the *ex-ante* determined endogenous variable vector. For this purpose, the multivariate cointegration techniques proposed by Søren Johansen (1988) and Johansen and Juselius (1990) are used. To briefly explain this method, let us assume a z_t vector of non-stationary n endogenous variables and model this vector as an unrestricted vector autoregression (VAR) involving up to k-lags of z_t:

$$z_t = \Pi_1 z_{t-1} + \Pi_2 z_{t-2} + ... + \Pi_k z_{t-k} + \varepsilon_t \tag{9}$$

where ε_t follows an i.i.d. process N(0, σ^2) and z is (nx1) and the Π_i is (nxn) matrix of parameters. Eq. 9 can be rewritten leading to a vector error correction (VEC) model of the form:

$$\Delta z_t = \Gamma_1 \Delta z_{t-1} + \Gamma_2 \Delta z_{t-2} + ... + \Gamma_{k-1} \Delta z_{t-k+1} + \Pi z_{t-k} + \varepsilon_t \tag{10}$$

where:

$$\Gamma_i = -I + \Pi_1 + ... + \Pi_i (i = 1, 2, ..., k-1) \tag{11}$$

and:

$$\Pi = \mathrm{I} - \Pi_1 - \Pi_2 - ... - \Pi_k \tag{12}$$

This specification of the system of variables carries on the knowledge of both the short- and the long-run adjustment to changes in z_t, via the estimates of Γ_i and Π. Following Richard Harris and Robert Sollis (2003), we can state that $\Pi = \alpha\beta'$. α measures the speed of adjustment coefficient of particular variables to a disturbance in the long-run equilibrium relationship and can be interpreted as a matrix of error correction terms. β is a matrix of long-run coefficients such that $\beta' z_{t-k}$ embedded in Eq. 10 represents up to (n-1) co-integration relations in the multivariate model which ensure that z_t converge to their long-run steady-state solutions.

For the lag length of the VAR model, we consider the sequential modified LR statistics employing Christopher Sims (1980) small sample modification which suggest the use of lag length 5. As a next step we estimate the long run co-integration relationships between the variables by using two likelihood test statistics known as maximum eigenvalue for the null hypothesis of r versus the alternative of r+1 co-integration relationships and trace for the null hypothesis of r co-integration relations against the alternative of n co-integration relations, for r = 0,1, ... , n-1 where n is the number of endogenous variables.

4. Results

The results of Johansen co-integration test are reported in Table 3 below using max-eigen and trace tests based on critical values taken from Michael Osterwald-Lenum (1992). Johansen (1992) and Harris and Sollis (2003) suggest the need to test the joint hypothesis of both the rank order and the deterministic components. In the case of a cointegration analysis, the limit distribution depends on the actual (true) number of the co-integration relations and also on the presence of a linear trend. Following Sastry G. Pantula (1989), they propose to identify the sub-hypotheses, which give different limit distributions, and construct a test statistic and a critical region for each of these sub-hypotheses. The hypothesis in question is only rejected if all sub-hypothesis are rejected. For this purpose, we restrict intercept and trend factors into the long-run variable space, but do not assume a quadratic deterministic trend lying in both the co-integration model and the dynamic vector error correction model. In line with such a rank determination procedure, we find that both LR tests tend to approve the existence of one potential stationary relationship in the long-term variable space as a cointegration vector.

Table 3. Rank Test Results

Hypothesized no. of vectors	Eigen-value	Trace	0.05 cv	Max-eigen	0.05 cv
None	0.504	98.50*	88.80	47.71*	38.33
At most 1	0.314	50.78	63.88	25.59	32.12
At most 2	0.153	25.20	42.92	11.27	25.82
At most 3	0.141	13.93	25.87	10.35	19.39
At most 4	0.051	3.58	12.52	3.58	12.52

* denotes the rejection of the unit root null hypothesis at the 5% level

Source: Authors' own calculations.

However, we must be somewhat more careful on this point, since it has just been possible that some structural breaks may be attributed to the rank order of the cointegration relationships especially for a country such as Turkey. Therefore, in order to test the existence of a cointegration relationship subject to structural breaks, we also employ the method suggested by Johansen, Rocco Masconi, and Bent Nielsen (2000), which can be used to specify up to two structural breaks either in levels or in levels and trend jointly. Here we tend to test the sensitivity of the rank results obtained above to some exogenous breaks in levels and trend jointly, allowing trend shift restricted to error correction term and level shift unrestricted in the model. We choose the exogenous break dates as 1994Q2 and 2000Q1 which coincide with the occurence of the macroeconomic crisis conditions within the Turkish economy. The results are reported in Table 4. Note that the critical values as well as the *p*-values are now taken from the Johansen trace tests and are obtained by computing the respective response surface estimates. Of course, an alternative method might be the estimation procedures suggested by Allan W. Gregory and Bruce E. Hansen (1996) which allow an endogenous break in the co-integration test. However, since the two enormous economic crises have been observed highly explicit as a diversification date in the data by ourselves, we chose the method of Johansen, Masconi, and Nielsen (2000) to apply to the Turkish data. In Table 4, we see that the null hypothesis of one co-integration vector cannot be rejected under the acceptance of two exogenous structural breaks attributed to the macroeconomic crisis conditions in the Turkish economy.

Table 4. Rank Test Results with Exogenous Breaks

Restricted Dummies 1994Q2 and 2001Q2
Trend and Intercept Included
Response Surface Computed

Hypothesized No. of vectors	LR	95%
None	132.85*	111.97
At most 1	80.14	82.95
At most 2	52.02	57.81
At most 3	29.25	36.41
At most 4	11.17	18.32

LR represent the relevant likelihood ratio test
* denotes the rejection of the unit root null hypothesis at the 5% level

Source: Authors' own calculations.

As a next step, we examined whether the relevant cointegration vector can give support to our *a priori* model expectations. For this purpose, the estimation results have been presented in Table 5.

We find that the first cointegration vector with the largest eigenvalue indeed satisfies our model consideration running from Eq. 1 to Eq. 4. All the explanatory factors have a positive significant impact on the Turkish inflation. Of all these, the most significant one is the nominal exchange rate depreciation carrying the largest coefficient in value. We are unable to reject the homogeneity restriction of the exchange rate changes to the foreign price changes within the cointegration relationship. We also cannot reject the linear homogeneity of the sum of all the explanatory factors as to the domestic inflation. In addition, we find a negative and significant normalized trend value, which explicitly reflects the downward trend in the changes of the consumer prices inside the period. If we explicitly write down in Eq. 13 this final cointegration equation, implying linear homogeneity of the sum of the coefficients to the domestic consumer price inflation (standard errors in parentheses):

Table 5. Estimation Results

Unrestricted Co-integration Coefficients

p_t	ulc_t	e_t	p_t^{for}	p_t^{ws}	$trend$
65.08	-6.234	-32.68	-20.44	-19.07	0.084
-70.36	22.21	14.75	-150.8	31.50	-0.022
-26.70	8.983	-2.052	182.8	36.30	0.142
-30.96	2.533	6.355	-70.13	24.69	0.019
-30.99	-3.104	-1.682	-62.84	28.64	-0.110

Unrestricted Adjustment Coefficients (D's are the difference operators)

$\Delta(p_t)$	-0.020	-0.001	-0.004	-0.002	0.002
$\Delta(ulc_t)$	0.002	-0.023	0.006	0.005	0.005
$\Delta(e_t)$	-0.028	-0.020	-0.011	-0.017	-0.001
$\Delta(p_t^{for})$	0.001	0.001	-0.001	0.001	-0.001
$\Delta(p_t^{ws})$	0.027	-0.005	-0.007	-0.001	-0.001

1 Normalized Co-integration Equation (standard errors in parentheses)

p_t	ulc_t	e_t	p_t^{for}	p_t^{ws}	$trend$
1.000	-0.096	-0.502	-0.314	-0.293	0.0012
	(0.036)	(0.055)	(0.061)	(0.082)	(0.0004)

Weak Exogeneity Test Results ~ $\chi^2(1)$ distribution

p_t	ulc_t	e_t	p_t^{for}	p_t^{ws}
21.88	0.069	6.833	0.864	20.20

Multivariate Statistics for Testing Stationarity ~ $\chi^2(4)$ distribution

p_t	ulc_t	e_t	p_t^{for}	p_t^{ws}
38.14	36.59	38.81	35.87	37.52

b(1,1)=1, b(1,3)=b(1,4)	$\chi^2(1)$=0.071 (prob. 0.790)
b(1,1)=1, b(1,2)+b(1,3)+b(1,4)+b(1,5)=-1	$\chi^2(1)$=0.086 (prob. 0.769)

VEC Residual Serial Correlation LM Test

Lag 4	LM-Stat 25.217	prob. 0.450

$$\beta' z_t = p_t - \underset{(0.034)}{0.102} ulc_t - \underset{(0.052)}{0.491} e_t - \underset{(0.039)}{0.112} p_t^{for} - \underset{(0.076)}{0.295} p_t^{ws} + \underset{(0.000)}{0.001} \text{trend} - 0.135 \quad (13)$$

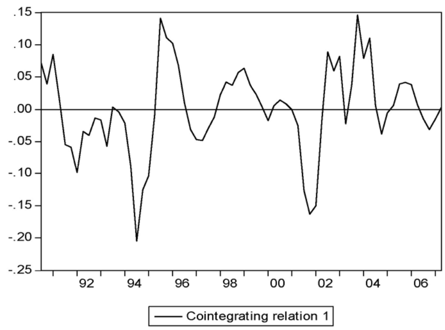
Source: Authors' own calculations.

As to the weak exogeneity characteristic of the variables, we are able to reject the null hypothesis for the consumer price inflation, nominal exchange rate and producer price inflation, but not for the relevant wage and foreign price data. These results should not be counted surprising since the wage data are mainly affected by the labor market conditions. Thus, even though the course of the price of labor can be considered one of the main determinants of the consumer price inflation, no feedback effect may be observed from the consumer inflation to the labor market data. Likewise, the weak exogeneity of the foreign inflation can normally be expected by the researchers due to the small open economy characteristic of the Turkish economy when compared with the US economy. Finally, multivariate statistics for testing stationarity are in line with the univariate unit root test results obtained above in the sense that no variable alone can represent a stationary relationship in the co-integration vector. Below in Fig. 2 is shown that the estimated relationship has really a stationary characteristic.

Having established the long-run cointegration model, we report the dynamic single equation parsimonious vector error correction model using both a reduced form model with the econometrically meaningful variables shown and the estimated error correction term (*EC*) produced in the cointegration relationship. Since all the variables in the model are now of a stationary form, statistical inferences using standard *t*- tests are valid. The results are reported in Eq. 14 *t*-stats are given in parentheses below the coefficients.

It is essential for maintaining the long run equilibrium conditions to reduce the existing disequilibrium in time. We find that the deviations from the long-run path of the cointegration data are corrected by about 69% within one period in a way indicating a highly quick adjustment process to the long-run equilibrium relationship. Economic theory is rarely interested in the short-run characteristics of the variables, but generally pays attention to the long run behavior of the variables. However, when we look at the estimated coefficients, we can notice that a nearly one-to-one positive effect from the changes in the two-period lagged inflation predominates within the parsinomious error correction model. This is an explicit indicator of the inertial nature of the changes in inflation for the Turkish economy. The net effect of the changes in the exchange rate growth on the changes in the domestic inflation is positive. The unit labor costs also have a similar characteristic. We find that there seems to be a highly strong positive total impact of the changes in the foreign consumer price inflation on the domestic inflation changes. Thus, all these reveal that the cost-push factors tend to determine the course of the domestic price changes in the short run, as well. But the behavior of the producer price inflation on the consumer price inflation turns out to be negative. On this point, we tend to neglect this anomaly as to our expectitons in the short run, since the all the other model properties give us a significant knowledge to explain both the short- and the long-run properties of the Turkish inflation.

Further, the parsimonious model has good diagnostics (probs in parentheses). We observe no serial correlation problem according to the Breusch-Godfrey (BG) test results. There exists no heteroskedasticity problem through the White tests, no residual non-normality problem through the Jarque-Bera (JB) statistics and no model misspecification problem through the RESET test.



Source: Authors' own calculations.

Figure 3 The Graph of the Co-integration Relationship

$$
\begin{array}{l}
D(p_t) = \underset{(-1.50)}{-0.01} - \underset{(-4.92)}{0.69EC_{t-1}} + \underset{(2.71)}{0.96D(p_{t-2})} + \underset{(2.91)}{0.14D(ulc_{t-2})} + \underset{(4.99)}{0.46D(e_{t-1})} - \\
\underset{(-5.18)}{0.22D(e_{t-4})} + \underset{(2.05)}{0.15D(e_{t-5})} + \underset{(2.05)}{2.56D(p_{t-2}^{for})} + \underset{(2.13)}{1.92D(p_{t-5}^{for})} - \underset{(-4.13)}{0.77D(p_{t-1}^{ws})} - \\
\underset{(-2.88)}{0.80D(p_{t-2}^{ws})} - \underset{(-2.36)}{0.34D(p_{t-5}^{ws})}
\end{array}
\tag{14}
$$

Adj.R^2=0.59, BG AR(1)=0.37 (0.54), BG AR(4)=1.21 (0.32), White=0.57 (0.84), JB=1.89 (0.39), RESET=0.27 (0.60)

5. Conclusions and Suggestions for Future Research

A chronic inflationary framework is one of the main properties that identifies the course of the Turkish business cycles for the last 30-years period and constitutes an important benchmark for economic agents in constructing their expectations. We observe in the paper that the data of the post-1980 period indicate that inflation tends to mainly be characterized with the realizations of self-peculiar characteristics as to the sub-periods, but has never been decreased to the single-digit levels till the mid-2000s.

In this paper, our contribution to the existing literature is to empirically examine the appropriateness of a cost-push model of the aggregate price-setting in the economy. To this end, a simple economic model has been developed and then tested in the light of some contemporaneous time series estimation techniques. Our results employing the multivariate cointegration methodology of the same order integrated variables reveal that all the explanatory factors *a priori* modeled have a positive impact on the inflation. We find that the most significant component contributing to the inflation is the nominal exchange rate depreciation carrying the largest coefficient in value. We also cannot reject the linear homogeneity of the sum of all the explanatory factors as to the domestic inflation. We must specify that our results generally give support to the literature cited in the paper in the sense that the cost-push factors, especially the exchange rate depreciations, indeed have a significant explanatory power over the Turkish consumer price inflation.

All these results suggest that both economic agents and policy makers should take account of the developments in the cost-based factors in the economy when they construct their decisions as to the future course of the price changes. Otherwise, an incomplete and possibly mistaken economic decision processes related to the future expectations could result in undesirable outcomes for both individuals and policy authorities. Of course, additional research and future papers considering more detailed investigation of the relationships extracted in this study would be complementary to our paper, so as to see the validity of the estimation results. Furthermore, papers relating the macro-level pricing behavior to the main characteristics of the business cycles should be constructed to examine the consistency of the results obtained in this paper with the cyclical properies of the Turkish economy.

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