Does the Law of One Price Hold
in a High-Inflation Environment?
A Tale of Two Cities in Turkey[§]

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Abstract

This study addresses price convergence in two cities in Turkey (Istanbul and Ankara) using the annual data over the three quarters of the 20th century (1922–1998), characterized by prevailing high inflation rates for most of the period. In contrast to the rest of the literature addressing convergence in price levels with a typical result of extremely slow convergence rates at best, we argue that convergence is much easier detected in growth rates rather than levels of prices. We suggest using the bounds testing procedure of Pesaran et al. (2001) for this purpose. We find a clear-cut evidence on the existence of a common driving force behind inflation dynamics in Istanbul and Ankara—a finding that is in contrast with the results typically reported in related literature.

Keywords: Price convergence, Bounds testing procedure, Turkey

JEL code: C22, C32, C52, E31.

1 Introduction

In this paper we suggest a novel approach to testing for price convergence in different geographical locations—a topic that attracted a lot of attention in the economic literature (e.g., see Sonora, 2009, 2008, 2005; Busetti, Forni, Harvey, and Venditti, 2007; Cecchetti, Mark, and Sonora, 2002; Goldberg and Knetter, 1997; Parsley and Wei, 1996; Rogoff, 1996; Froot, Kim, and Rogoff, 1995; Isard, 1977, among many others). An intuitively appealing argument behind purchasing power parity (PPP) that national price levels, once converted to a common currency, should be the same typically finds a little empirical support prompting Rogoff (1996) to introduce the purchasing power parity puzzle in his seminal contribution. Although such an extremely slow convergence to PPP observed at the international level could be at least partially explained by prevailing trade barriers, exchange rate volatility, sticky nominal wages and prices, it is even more puzzling that a similar conclusion has been typically achieved also at the intranational level characterized by a common economic space involving a common currency and the absence of trade barriers.

For example, for Turkey Özcicek (2001) investigated the price convergence among 19 Turkish provinces for the period covering 1994M1 till 2003M12. Özcicek (2001) finds no evidence of price convergence on the basis of univariate and panel unit root tests.

In sequel, we argue that a typical finding of puzzlingly slow convergence—or even its absence—between price levels in different geographical markets could be explained by an improper modeling of the underlying time series. Most of the applied research investigating the law of one price typically fails to account for the presence of structural breaks leading to spurious finding of high persistence (non-stationarity) either in prices and/or, more importantly, in relative prices, which as implied by PPP should be stationary, see Perron (1989) for consequences of unmodeled structural breaks on unit-root test outcomes.

We reconsider the issue of price convergence in Turkey and hence revise the findings of Özcicek (2001) in several ways. First, we use the long time series data on prices collected at the annual frequency for two Turkish cities (Ankara and Istanbul) for the period covering 1922 till 1998. The period is much longer than that in Özcicek (2001) allowing us to focus on a truly long-run relationship between these two price indices. Several periods of very high inflation rates makes our exercise more interesting as it allows us investigating whether the law of one price held also in such high-inflation environment.

Second, instead of looking for convergence among price indices we opt for testing whether the law of one price held in growth rates of prices or inflation. Thus, our approach of testing for price convergence is consistent with the relative PPP described in Frenkel (1978, 1981) where in the international context changes in exchange rate are related to changes in inflation differential. In the intranational context with a common currency, the version of relative PPP reduces to testing whether there exist a long-run relationship in the form of inflation differential. Existence of such a long-run relationship would imply that inflation dynamics is governed by a common factor whose presence rules out persistent deviations in inflation in the long run, implying that markets are integrated and the law of one price holds in a relative form.

Third, by focusing on inflation rather than on price levels greatly facilitates treatment of structural breaks in time series under scrutiny. These structural breaks are very likely to occur in our data set as it

covers more than three quarters of the 20th century. We accommodate these structural breaks by means of impulse dummies. Observe that what appears as a spike in inflation translates itself to a step-wise shift in the associated price levels data, i.e., outliers of a different magnitude observed in inflation rates result in persistent deviations of price levels, which are much more difficult to account for and that are likely to explain the puzzling results of extremely slow (if any) convergence between price levels that is typically reported in the relevant literature.

Last but not least, we propose to address the existence of the law of one price in two distant markets by testing for existence of a long-run relationship between inflation rates by employing the bounds testing procedure of Pesaran, Shin, and Smith (2001). The advantage of using this procedure is that it can be applied in cases when regressors can be I(1), I(0), or mutually cointegrated. Furthermore, the procedure is based on an unrestricted error-correction model, which permits joint estimation of long- as well as short-run effects. As pointed out in Banerjee, Dolado, and Mestre (1998), joint estimation has better statistical properties than the two-step Engle-Granger procedure that pushes the short-run dynamics into the error term.

The use of the procedure of Pesaran, Shin, and Smith (2001) in the current context also can be supported by the fact that so far there is no uniform agreement in the literature on whether (logs of) prices should be modeled as an I(2) or I(1) process. This naturally has implications for the order of integration of growth rates of prices or inflation. In the former case inflation is also a unit-root process, whereas in the latter case inflation, correspondingly, should be modeled as an I(0) process. On the one hand, a common observation of a rather high persistence in inflation rates supports the former view reflected in Banerjee, Cockerell, and Russell (2001) and Juselius (1995), for example, where prices were explicitly assumed to be I(2) variables. However, treating inflation as a unit-root variable is at odds with economic models such as the sticky price model of Taylor (1979) or the Phillips curve model proposed in Calvo (1983), for instance. On the other hand, Hendry (2001) views the price levels data as integrated of order one (I(1)) with superimposed major breaks such that they mimic behavior of I(2) processes. As a consequence, inflation is modeled as an I(0) process with breaks.

In a recent paper Romero-Ávila and Usabiaga (2009) report a clear-cut evidence based on unit root tests carried out in a panel-data setting jointly accounting for cross-sectional dependence and for the presence of unknown number of breaks that inflation in selected 13 OECD countries can be considered as an I(0) process with structural breaks. This finding supports regime-wise stationarity of inflation.

Our main finding is that we find a clear-cut evidence concerning existence of a long-run relationship between inflation rates in Ankara and Istanbul supporting an intuitively appealing notion of a common driving force behind price dynamics in these two distant markets in Turkey. Our finding is in sharp contrast to that reported in Özcicek (2001). The fact that Özcicek (2001) addresses price convergence in price levels without proper accounting for the presence of structural breaks is a likely (but, perhaps, not the only) reason that explains the different conclusion from ours.

The rest of the paper is organized as follows. Section 2 contains the description of data and its sources. In Sections 3 and 4 we describe methodology applied and report estimation results. The final section concludes.

2 Data

All the data on price indices come from a single source Pamuk (2000) and represent a constructed Consumer Price Index (CPI) based on a comparable consumption basket containing both food and non-food items compiled from the following sources: data for the period from 1914 till 1937 is due to Pamuk (2000, p. 22–23 for Istanbul and p. 58–59 for Ankara), for the period from 1938 till 1987—the Undersecretariat of Turkish Treasury, and for the period since 1988—the Turkish Statistical Institute[1]. Due to the fact that there are no continuous data for Ankara during the period from 1914 till 1922, we restrict our estimation sample to start from 1922. According to Pamuk (2000), the same base year 1914 has been chosen for both price indices.

In sequel, we will denote the CPI for Ankara and Istanbul in levels as p^A and p^I and the corresponding logarithmic transformation of these two price indices as $\ln p^A$ and $\ln p^I$. Furthermore, the inflation in Ankara and Istanbul is denoted as $INF^A = \Delta \ln p^A$ and $INF^I = \Delta \ln p^I$, respectively. Both the price indices (in logs) as well as the corresponding inflation rates are displayed in Figure 1.

The overall impression from Figure 1 is that there is certain common features shared by the price indices in these two cities: a relatively stable prices during 1920s that even decrease during the pre-WWII period, a stepwise increase during the beginning of 1940s and then further stabilization in the late 1940s. Both price indices have tendency to increase since 1950s and, more importantly, this tendency became more pronounced since the late 1970s. The similar dynamics is also reflected in the pair of inflation time series: a moderate inflation rates during 1920s followed by deflation in the pre-WWII period, a spike in inflation rates during the beginning of 1940s, reflected in a stepwise shift in the corresponding price levels, a relatively stable inflation in the post-war period, which started to increase during 1950s, and then falling back to the post-war levels in 1960s. Both inflation rates display an upwards trending behavior since the early 1970s with occasional outbursts, e.g., in the late 1970s and in mid-1990s. For a comprehensive review of inflation experience in Turkey we direct an interested reader to Kibritçioğlu (2002).

The difference of log-price levels between Istanbul and Ankara are displayed in Figure 2; unsurprisingly, there are quite persistent deviations to be observed, implying that the tests for price convergence when applied to price levels directly are likely to find no evidence supporting it, see Özcicek (2001) for an example using the Turkish data. On the contrary, the cross-plot of inflation rates in Istanbul and Ankara presented in the right top panel suggests that there is a close relationship between inflation rates in these two cities. The two lower panels in the same figure present changes in inflation which look quite similar aside from the fact that changes inflation in Ankara have been more volatile in the 1980s than in Istanbul.

All in all, visual inspection suggests that the inflation rates in Ankara and Istanbul tend to move together, suggesting that these two markets are integrated and react similarly to common shocks. It remains to see whether this informal conclusion will be verified by application of statistical methods.

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[1] For the period until 1937 the consumption basket includes the following food items: wheat, flour, bread, rice, cooking margarine, olive oil, lamb meat, mutton meat, honey, sugar (cube and pulverized), coffee, chickpeas, lentils, onion, small onion, milk, eggs, tea, and the following non-food items: soap, coal, wood, nails, cloth, London cloth, velvet, satin (see Basvakalet, 1946; Bulutay et al., 1974). The item classification for construction of the CPI used by the Undersecretariat of Turkish Treasury and by the Turkish Statistical Institute since 1938 is as follows: foodstuffs, housing, household expenditure, clothing, medical care, transportation, culture and entertainment.

3 Methodology

In this section we briefly describe the bounds testing procedure of Pesaran et al. (2001) that we apply for testing whether there exists a long-run relationship between inflation in Ankara and Istanbul. The bounds testing approach has broad applicability since the regressors can be either I(1), I(0), or mutually cointegrated, which is of a particular importance in our case given inconclusive results on the order of integration of inflation. Secondly, as discussed above the structural breaks are much easier accommodated into the model estimated for inflation rather than for price levels.

As a starting point for implementing the bounds testing procedure we assume that inflation in Istanbul INF^I and inflation in Ankara INF^A are related according to a vector autoregressive (VAR) model of order p that is further reduced to the following conditional error correction model (ECM),

$$\Delta INF_t^I = \alpha + \theta_0 INF_{t-1}^I + \theta_1 INF_{t-1}^A + \sum_{i=1}^{p-1} \lambda_i \Delta INF_{t-i}^I + \sum_{i=0}^{p-1} \beta_i \Delta INF_{t-i}^A + \omega' D_t + \varepsilon_t. \tag{1}$$

The lagged values of INF^I and INF^A form a long-run relationship. The deterministic terms such as a constant, and dummy variables are denoted by α and D_t, respectively. The short-run dynamics is captured by means of lagged values of ΔINF_t^I and current and lagged values of ΔINF_t^A. The long-run relationship between inflation in Istanbul and in Ankara is given by the following vector $(1, -\theta_1/\theta_0)'$ (Banerjee et al., 1998). Observe that if inflation in these two cities react homogenously to shocks then the vector of interest reduces to $(1, -1)'$.

The bounds testing procedure uses the conventional F-test for testing the null hypothesis $H_0 : \theta_0 = \theta_1 = 0$. Note that this statistic has a non-standard distribution which depends upon (i) the order of integration of the regressors; (ii) the number of regressors; (iii) the set of deterministic terms included in the model; and (iv) sample size. Pesaran et al. (2001) provide the set of asymptotic critical values. We, however, in the hypothesis testing rely on the critical values simulated in Narayan (2005) for a sample size comparable to ours.

There are two sets of critical values. The first set gives the lower bound, applicable when all regressors are I(0). The second gives the upper bound, applicable when all regressors are I(1). If the calculated F-statistic falls below the lower bound, the null hypothesis of no relationship between inflation in both cities cannot be rejected. Conversely, if the F-statistic exceeds the upper bound, the null hypothesis of no long-run relationship is rejected. As noted above, these critical bounds can be applied irrespective of the order of integration of the regressors. Finally, if the F-statistic falls within the critical bounds, the order of integration of the variables must be established in order to obtain conclusive inference.

In addition, as argued in Pesaran et al. (2001, footnote 17, p. 307) the asymptotic theory and the associated critical values must not be modified provided that fraction of observations for which one uses dummy variables tends to zero as the total sample increases.

4 Results

Table 1 presents the results of the lag order p selection procedure for Equation (1). The information criteria (Akaike, AIC and Schwarz, SIC) as well as the Lagrange Multiplier statistic testing for remaining autocorrelation up to the first and second orders in regression residuals are reported.

Both information criteria—AIC as well as SIC—select $p = 1$. For all considered values of p, there is no evidence of remaining autocorrelation in the regression residuals. Given the results from the selection criteria and the evidence of no residual autocorrelation regardless of the value of p, the model with $p = 1$ is preferred. Observe that in order to account for the presence of outliers corresponding to the periods of unusually large discrepancies between inflation rates in Istanbul and Ankara the following impulse dummies (DYY_t) have been included in the test regression: $D27_t$, $D29_t$ and $D44_t$ each corresponding to the year $19YY$[2].

The corresponding F-test statistic for the joint null hypothesis $H_0 : \theta_0 = \theta_1 = 0$ using the finite-sample critical values simulated in Narayan (2005) for $T = 75$ corresponding to case III in Pesaran et al. (2001), i.e., with unrestricted constant and no linear deterministic trend, is reported in the last column of Table 1. As seen, the null hypothesis of no long-run relationship between inflation in Istanbul and Ankara can be decisively rejected for $p = 1$ and $p = 2$ at the 1% significance level. For $p = 3$ the test statistic falls inside the bounds implying that without further pre-testing no definite conclusions on the existence of a long-run relationship between inflation in these two cities can be reached. This indefinte results is, however, due to fact that the model with $p = 3$ is clearly over-parametrized as indicated by the values of the information criteria.

Having established the existence of a long-run relationship between inflation in Istanbul and Ankara we can estimate the coefficients of interest. Starting with the error-correction model corresponding to $p = 2$ and after deleting the insignificant augmentation lags we arrive at the following parsimonious model[3]:

$$
\begin{aligned}
\Delta INF_t^I \;=\; &\underset{(0.0065)}{0.0117} \;+\; \underset{(0.053)}{0.808}\ \Delta INF_t^A \;-\; \underset{(0.043)}{0.252}\ D27_t \\[2mm]
&+\; \underset{(0.043)}{0.156}\ D29_t \;-\; \underset{(0.046)}{0.156}\ D44_t \;-\; \underset{(0.049)}{0.118}\ D81_t \\[2mm]
&-\; \underset{(0.092)}{1.042}\ INF_{t-1}^I \;+\; \underset{(0.093)}{1.040}\ INF_{t-1}^A
\end{aligned}
$$

$R^2 = 0.873$, $\;F_{(7,67)} = 65.93[0.000]$, $\;T = 75$, $\;F_{(2,65)}^{AR(1-2)} = 0.013[0.986]$,

$F_{(1,65)}^{ARCH(1)} = 0.037[0.981]$, $\;\chi_{(2)}^{Norm} = 0.073[0.964]$, $\;F_{(9,58)}^{Hetero} = 0.776[0.651]$,

$F_{(1,66)}^{RESET} = 0.070[0.792]$,

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[2]The outliers have been identified as those residuals exceeding regression standard error by factor two in the estimated regression (1) with $p = 1$ without intervention dummies. We also identified a recording mistake (see Pamuk, 2000, p. 59) for the price index in Ankara: the figure 13010 appears both in 1984 and 1985. In private correspondence, S. Pamuk provided us with the correct figure for 1984, 8652, which we use in the current analysis.

[3]Observe that in order to account for a moderate outlier in 1981 we inserted an additional impulse dummy for this year in our regression.

where standard errors are reported in parentheses and error probabilities—in brackets. The model above passes the standard specification tests such as tests of no residual autocorrelation, of no residual ARCH effects, of residual normality, of no residual heteroskedasticity, and the RESET test for functional form misspecification. The overall impression is that this parsimonious model delivers a satisfactory fit to data considering that the period under investigation that stretches over three quarters of the 20th century characterized by the Second World War, several domestic political and economic crises, two international oil crises and major legislative and technological changes.

The estimated model allows us to compare the coefficients belonging to the lagged inflation variables. These coefficients are of a similar absolute magnitude with the implied long-run vector of $(1, -0.998)'$ such that one can safely impose a homogeneity restriction $\theta_0 = -\theta_1$, i.e., the long-run relationship vector between inflation in Istanbul and Ankara is $(1, -1)'$. The restricted error-correction model is reported below:

$$
\begin{aligned}
\Delta INF_t^I \;=\; & \underset{(0.005)}{0.0113} \;+\; \underset{(0.053)}{0.808}\ \Delta INF_t^A \;-\; \underset{(0.043)}{0.251}\ D27_t \\[2mm]
& +\; \underset{(0.042)}{0.156}\ D29_t \;-\; \underset{(0.045)}{0.156}\ D44_t \;-\; \underset{(0.047)}{0.119}\ D81_t \\[2mm]
& -\; \underset{(0.091)}{1.042}\ (INF_{t-1}^I - INF_{t-1}^A)
\end{aligned}
$$

$$
R^2 = 0.873, \quad F_{(6,68)} = 78.06[0.000], \quad T = 75, \quad F_{(2,66)}^{AR(1-2)} = 0.013[0.987],
$$
$$
F_{(1,66)}^{ARCH(1)} = 0.020[0.887], \quad \chi_{(2)}^{Norm} = 0.044[0.978], \quad F_{(8,59)}^{Hetero} = 0.810[0.597],
$$
$$
F_{(1,67)}^{RESET} = 0.058[0.811],
$$

Imposing the homogeneity long-run restriction brings no noticeable changes in the estimated coefficients. As before all retained coefficients including those belonging to the impulse dummies are estimated with a high degree of precision. These encouraging results are also supported by the close match between actual and the fitted values displayed in the left top panel of Figure 3; the corresponding cross-plot is displayed in the right top panel. The estimated regression residuals and their autocorrelation function up to the ninth order are reported on the same figure in the left and right bottom panels, respectively. Finally, the results of the Chow tests for recursive stability and the recursive estimates of the model parameters are shown in Figures 4 and 5, respectively. In Figure 4 the values of the one-step, breakpoint, and forecast Chow test statistics are scaled by their respective 1% critical values (Doornik and Hendry, 2001). None of the tests show any sign of model instability.

5   Conclusion

We suggest a novel approach to testing whether the law of one price holds in the long run between different geographically distant markets. Inspired by the relative PPP hypothesis (see Frenkel, 1978, 1981) we suggest to conduct testing for market integration using growth rates rather than levels of prices.

To this end, we propose using the bounds testing procedure of Pesaran et al. (2001) which can be used in

situations when there is no consensus in the literature on the order of integration of the modeled variables. In particular, it can be used in situations when regressors can be either I(0), I(1) and/or mutually cointegrated. Another advantage of our approach is that the presence of structural breaks can be much easier addressed when testing for the presence of a long-run relationship between inflation rates rather than price levels in different markets. Intuitively, an extraordinary large spike in inflation rate in one location that is unmatched in magnitude in inflation rate in another location results in an unusually large residual for a single period when modeling the long-run relationship between inflation. In this case its influence can be captured by an impulse dummy inserted in this particular period. On the contrary, when testing for convergence in levels of prices such a spike translates into a persistent deviation of one price level from another that has to be captured by a stepwise shift or an intercept correction which is more difficult to implement in practice. As the previous research shows failure to accommodate structural breaks when testing for convergence between price levels typically results in puzzling results of extremely slow (if any) convergence (see Özcicek, 2001, for an example testing for price convergence in Turkey).

We illustrate our approach by testing whether the law of one price holds between Istanbul and Ankara using the inflation time series covering three quarters of the 20th century, from 1922 until 1998. This period is characterized by the Second World War, several domestic political and economic crises, two international oil crises and major legislative and technological changes. Needless to say, that during most of the period under scrutiny very high inflation rates prevailed in Turkey. Despite all this we find a clear-cut evidence on the existence of a common driving force behind inflation dynamics in Istanbul and Ankara—a finding that is intuitively appealing from the point of view of economic theory. This finding signifies that the economies of these two cities have been well integrated not only during normal times but also during periods of high inflation.

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Table 1: Lag order selection, 1926-1998

p	AIC	SIC	AR(1)	AR(2)	$F^{III}_{H_0:\theta_0=\theta_1=0}$
1	**-6.256**	**-6.036**	0.772	0.955	57.825***
2	-6.206	-5.924	0.508	0.804	18.653***
3	-6.213	-5.868	0.798	0.522	7.591$^+$

Notes: p is the lag order of the underlying VAR model for the conditional ECM, see Equation (1). AIC and SIC are the Akaike and Schwarz Information Criteria, respectively. AR(1) and AR(2) are the p-values of the Lagrange Multiplier test statistics for testing for residual autocorrelation of orders up to one and two, respectively. Bold entries indicate the lag order for which the respective values of information criteria are minimized.

$F^{III}_{H_0:\theta_0=\theta_1=0}$ denotes the F-test statistic for the null hypothesis $H_0 : \theta_0 = \theta_1 = 0$ using the finite-sample critical values reported in Narayan (2005) for $T = 75$ corresponding to case III in Pesaran et al. (2001), i.e., with unrestricted constant and no linear deterministic trend. '***' indicates that the null hypothesis of interest can be rejected at the 1% significance level. '$^+$' indicates that the test statistic falls inside the bounds (see Narayan, 2005, p. 1988) for T=75.
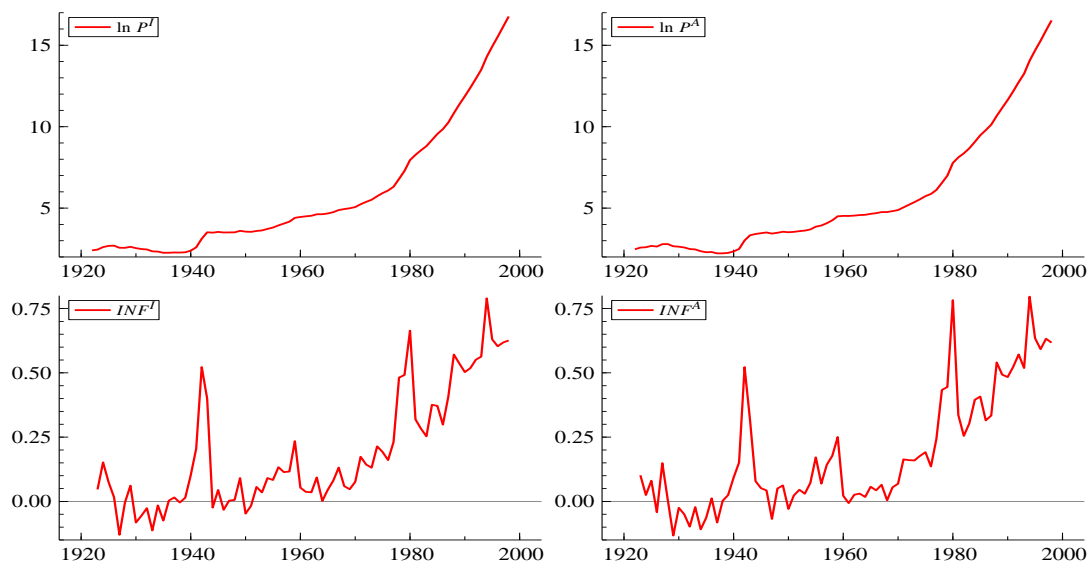
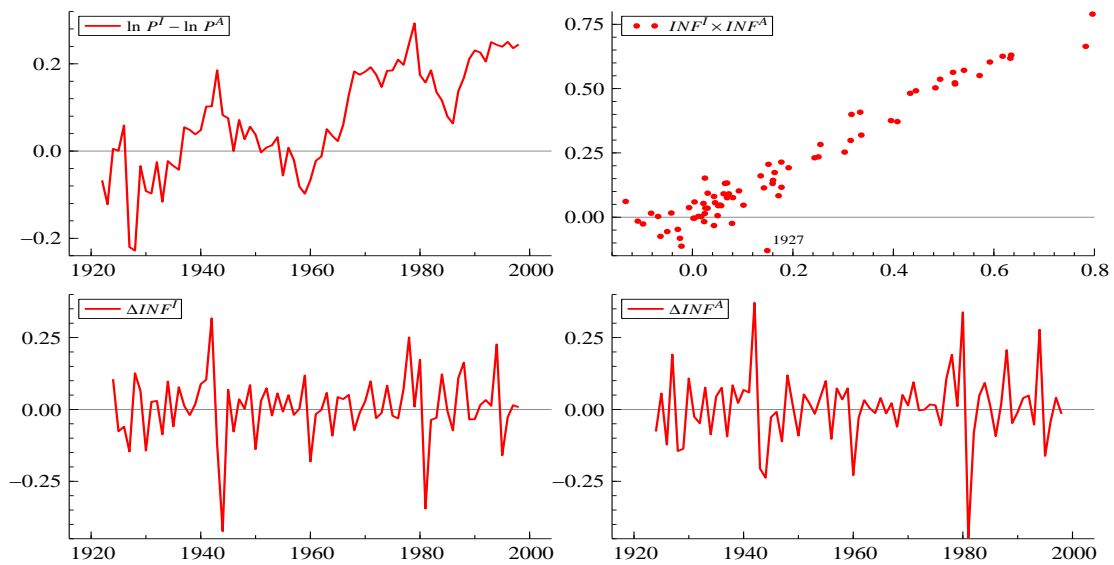Figure 1: Actual data: Price level (in logs) and inflation in Istanbul and Ankara

Figure 2: Actual data: Log-price differential $(\ln p^I - \ln p^A)$; cross-plot of inflation in Istanbul and Ankara INF^I and INF^A); changes in inflation ΔINF^I and ΔINF^A in Istanbul and Ankara
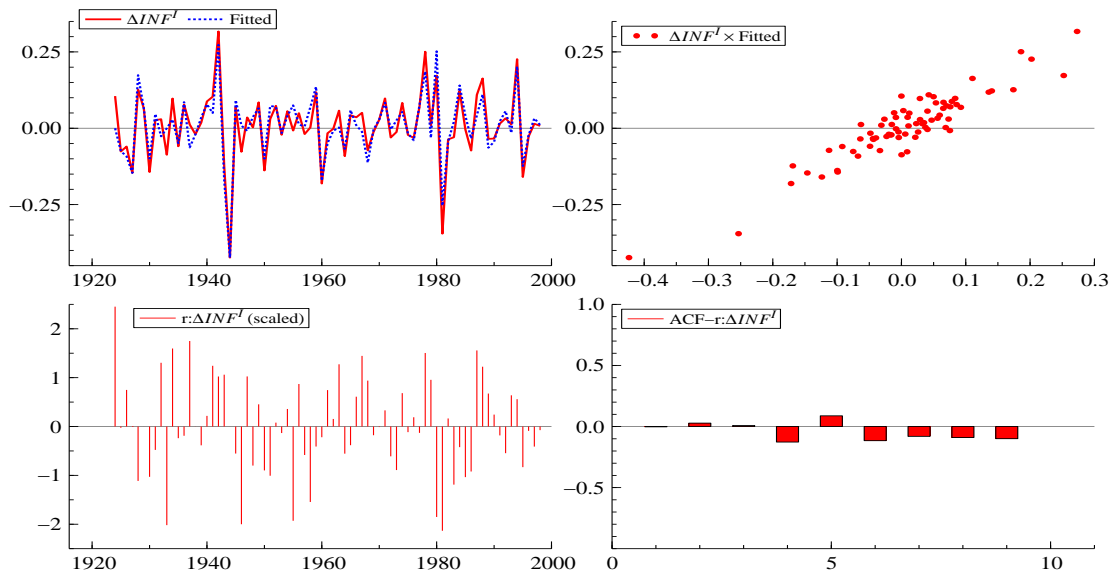
Figure 3: Actual and fitted values; Cross-plot of actual and fitted values; Regression residuals (r:ΔINF^I); Autocorrelation function of regression residuals (ACF-r:ΔINF^I)
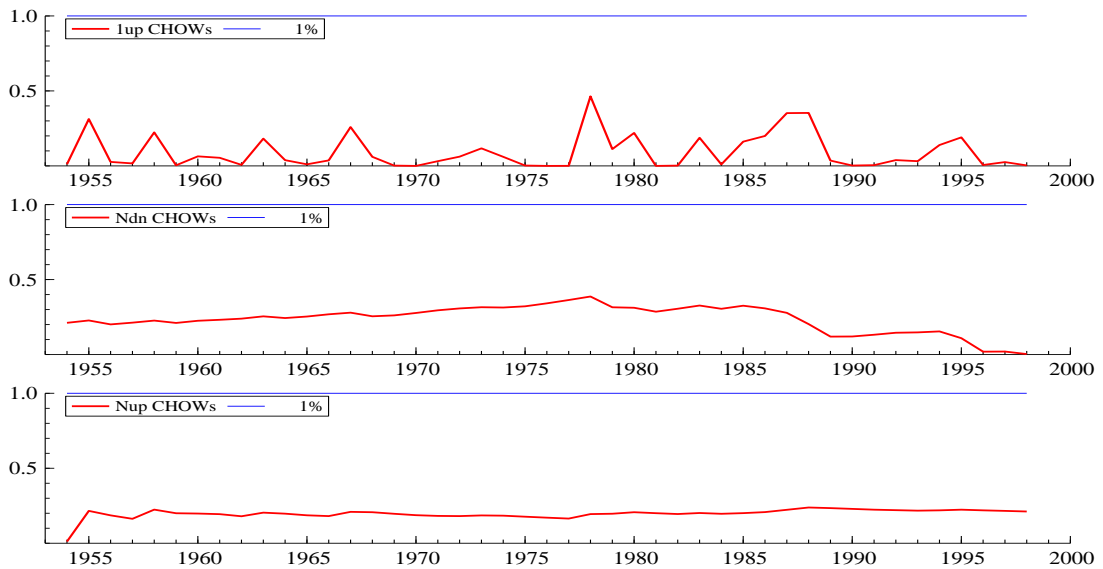

Figure 4: Recursive stability 1-step, breakpoint, and forecast Chow test statistics scaled by their respective 1% critical values.
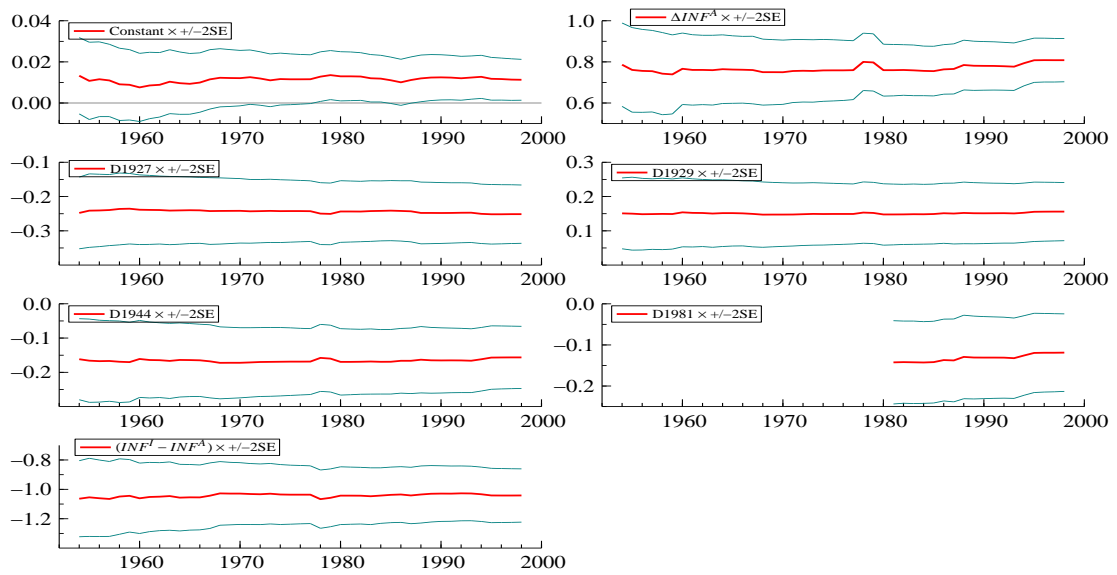
Figure 5: Recursive estimates of model coefficients