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Trend Breaks and Nonstationarity in the Yugoslav Black Market for Dollars, 1974-1987

by

John W. Dawson*, Steven W. Millsaps, and Mark C. Strazicich

Department of Economics
Appalachian State University
Boone, NC 28608-2051, USA

Abstract

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Keywords: Black market exchange rate, structural breaks

JEL Classification: C22, F31

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*Corresponding author: Phone 828-262-6128; email address: dawsonjw@appstate.edu.

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Abstract

We estimate a model of the black market premium for dollars in Yugoslavia from 1974-1987. Unlike previous applications of the model, our analysis addresses nonstationarity in the underlying data by allowing for trend breaks. Endogenous structural break tests indicate the presence of breaks closely associated with the death of Tito and changes in laws affecting the operation of the black market. After accounting for these breaks, we find strong support for the underlying model. In addition, we find evidence consistent with the era of increased government involvement in the black market leading to greater volatility of the premium following regime change.

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I. Introduction

This paper applies the model of Dornbusch *et al* (1983) to the black market for dollars in Yugoslavia from 1974-1987. The model predicts that the level of the black-market premium--the percentage excess of the black market price of dollars over the official exchange rate set by the monetary authority--is affected by the official real exchange rate, depreciation-adjusted interest rate differentials, and seasonal factors associated with tourism and import/export smuggling in a country with an inconvertible currency and an officially determined exchange rate. While existing empirical studies generally lend support to the model, our application to Yugoslavia involves an empirical approach not previously considered in the literature. Specifically, we address nonstationarity in the levels of the variables by allowing for structural changes in the black market exchange rate model.[1] The Yugoslav example provides an interesting demonstration of this approach because regime change associated with the death of Tito in 1980 is followed shortly thereafter by legal changes relating to the operation of the black market and an era of considerable government intervention in the black market, all of which are potential sources of structural change. This setting also allows us to address the impact of government involvement in the black market--in particular, in explaining the marked increase in the volatility of the black-market premium during the early 1980s.[2]

The paper is organized as follows. Section II provides a brief account of black market

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[1]Applications of the Dornbusch *et al* model include Phylaktis (1992), Shachmurove (1999), and Bahmani-Oskooee and Goswami (2005), among others. These studies and many others do not address the issue of potential nonstationarity. Other studies use a cointegration approach to address nonstationarity in the data. See, for example, Moore and Phylaktis (2000) and Diamandis and Drakos (2005). No studies that we are aware of allow for the possibility of structural changes in the model.

[2]Shachmurove (1999) includes Yugoslavia in a study of the black-market premium in a panel of developing countries. However, the time dimension of the study is 1985-1989, which is too late to allow any consideration of the issues noted here.

activities in Yugoslavia. Section III describes the model and the time series properties of the underlying data. Section IV estimates the model and discusses the results. Section V discusses the volatility of the black-market premium, and section VI concludes. An appendix provides a review of the trend break tests used in the analysis.

II. A Brief History of the Black Market in Yugoslavia

Black-market activity in Yugoslavia has persisted since the dinar was issued in 1931. The government tolerated the black market as long as residents obtained only medium-sized amounts of foreign currency for travel. However, the regime frowned upon large-size transfers of flight capital through the black market by syndicates and import/export smugglers. Numerous edicts regulated import and export operations. Although technically illegal, the black market was readily accessible to Yugoslavs and it functioned openly for years as a parallel operation to the government monetary authority.

In 1982, just two years after the death of Tito, Yugoslavia began a long scramble for foreign exchange primarily to pay off loans to the IMF and western creditors. The nation was starved for foreign exchange. In May of 1982, a new law empowered the state the right to take 75% of hard currency exchange held by any enterprise, including private citizens. Following a 16.7% downgrade of the dinar in October of 1982, the spreading black market received governmental attention. Stiff fines were imposed on residents dealing in the currency black market. Foreign tourists were given dinar-denominated checks instead of dinar banknotes upon conversion of hard currency. As late as 1986, the government announced a new Law of Foreign Exchange Operations that allowed considerable government regulation of the black market.[3]

[3]For a complete discussion of the evolution of government intervention in the black market in Yugoslavia, see International Currency Analysis, Inc., *World Currency Yearbook* (1989 and selected years) and Shaplen (1984).

In summary, three important dates to remember for our discussion of trend breaks in the next section are: (1) the May, 1980 death of Tito; (2) the May, 1982 law empowering state activity in the black market; and (3) the 1986 Law of Foreign Exchange Operations.

III. Model and Data

III.A. The Model

The Dornbusch *et al* (1983) model of the black-market premium suggests the following regression equation:

$$PREM_t = b_0 + b_1 RX_t + b_2 DI_t + \boldsymbol{b'MONTH} + u_t, \qquad (1)$$

where *PREM* is the black-market premium, *RX* is the official real (dinar/dollar) exchange rate, *DI* is the depreciation-adjusted interest differential between the dollar and dinar, ***MONTH*** = [*FEB*, ..., *DEC*]' is a vector of monthly dummy variables (January is omitted because a constant is included), $\boldsymbol{b} = [b_3,..., b_{13}]'$, and u is a normally and independently distributed error term. The model suggests that the premium on black-market dollars is negatively related to the official real exchange rate. This is because a real depreciation in the official exchange rate will lead to an increase in dollar inflows (an increase in the supply of black-market dollars) and put downward pressure on the premium.[4] Alternatively, the premium is positively related to the depreciation-adjusted interest differential, since a rise in the interest differential caused, say, by an increase in the nominal rate of interest on dollars, puts upward pressure on the premium, as dollars are now relatively more attractive and will cost more in terms of dinars. Thus, the expectation is that $b_1 < 0$ and $b_2 > 0$ in (1).

An important source of net inflow of dollars into the black market in Yugoslavia was due to tourism, especially in the summer months along the Adriatic coast from Split to Dubrovnik.

[4] An increase in *RX* denotes a real depreciation.

These months produced a seasonally high rate of dollar inflow into the black market giving rise to a seasonal accumulation of dollars and a seasonal decline in the premium. Accordingly, it is expected that the monthly dummies will capture the seasonal evolution in the premium that resembles this pattern.

III.B. The Data

Monthly data on Yugoslavia's black market exchange rate are taken from International Currency Analysis, Inc., *World Currency Yearbook* (various years). Other data are from the International Monetary Fund's *International Financial Statistics* (various years). The end-of-month nominal dinar-dollar exchange rate is used as the official exchange rate. The black market premium, *PREM*, is the percentage excess of the black market rate over the official rate. The official rate is multiplied by the ratio of the U.S. producer price index to the Yugoslav producer price index to calculate the official real exchange rate, *RX*. Short-term (3-month) nominal interest rates are used to calculate the depreciation-adjusted interest differential, *DI*, which is defined as $i_{\mathrm{US}} + d - i_{\mathrm{Yugo}}$, where i_{US} and i_{Yugo} are nominal monthly interest rates on dollars and dinars, respectively, and d is the rate of depreciation of the dinar in the black market.

The sample period for the analysis is 1974-1987. We choose to end our sample in 1987, just prior to the disintegration of the federal republic and subsequent period of hyperinflation, because the *DI* series begins to display extreme outliers as early as 1988. Furthermore, some of the newly formed states which emerged following the fall of Yugoslavia introduced their own currencies in the early 1990s, as the dinar entered a period of frequent devaluations. Since existing endogenous structural-break tests can identify at most two breaks, including a period as volatile as the post-1987 period in our sample would detract from the ability of these tests to identify the regime and legal changes of the early 1980s, which is the focus of this paper.

Figures 1-3 show the variables *PREM*, *RX*, and *DI*, respectively. Visual inspection of these plots suggests that all three variables are likely to be nonstationary, and that structural changes occur in both the level and volatility of these variables sometime after the early 1980s. Augmented Dickey-Fuller tests indicate that the null hypothesis of a unit root cannot be rejected for each of the three series.[5] Table 1 also suggests a change in behavior in the early 1980s, as the mean and standard deviation of each series is noticeably different before and after this period.[6]

III.C. Evidence on Trend Breaks in the Data

As noted above, conventional Dickey-Fuller tests imply that each of the series shown in Figures 1-3 are nonstationary processes. Recent developments in the time series econometrics literature suggest, however, that these tests fail to reject the null hypothesis of a unit root too often when the true data generating process is in fact trend stationary around a permanent break in the intercept and/or slope of the trend function. Thus, the structural change which was suggested in the discussion of Table 1 may be related to the finding that these series are nonstationary using conventional unit root tests.

Additional evidence on the presence of trend breaks in the data can be obtained by applying formal tests of structural change to the series *PREM*, *RX*, and *DI*. We use the two-break "minimum LM" test proposed by Lee and Strazicich (2003).[7] The results are summarized in Table 2. For all three series, the null hypothesis of a unit root is rejected in favor of the trend-break stationary alternative. Thus, in using these series to estimate the model in (1), de-trending

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[5]Results for these tests are available upon request from the authors.

[6]The "break date" used to construct Table 1 is October, 1982, based only on casual observation of the *PREM* series.

[7]A description of trend-break tests, in general, and the Lee and Strazicich (2003) test, in particular, is provided in the Appendix.

is appropriate to render the series stationary. The estimated trends for the *PREM*, *RX*, and *DI* series are shown in Figures 1, 2, and 3, respectively, along with the original series.

It is interesting to note that breaks are found in each of the series during the early 1980s, with several of the estimated break dates coinciding with the actual dates of important regime or legal changes in Yugoslavia. For example, the first identified break in the *RX* series is May, 1980, coinciding with the date of Tito's death. Similarly, the two breaks in the *PREM* series are estimated to be May, 1982 and June, 1986, corresponding to the dates of important laws affecting black market activity (as discussed in Section II). Thus, the empirical results are consistent with the hypothesis that regime change caused structural changes in these series.[8]

IV. Estimation and Results

We now turn to the estimation of the Dornbusch *et al* (1983) model of the black-market premium given by equation (1). As is well known in the econometrics literature, it is, in general, not appropriate to estimate a regression model using variables that are nonstationary. Following the discussion in the previous section, we utilize the identified breaks to de-trend the *PREM*, *RX*, and *DI* series and render them stationary.[9] We denote the resulting de-trended stationary series as $PREM^{det}$, RX^{det}, and DI^{det}, respectively.[10] Thus, the model to be estimated is:

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[8]Other commonly used one-break endogenous unit root tests, such as those proposed by Zivot and Andrews (1992) and Vogelsang and Perron (1998), also suggest breaks in the early 1980s for the *PREM* and *DI* series, but not the *RX* series. See the discussion in the Appendix for the advantages of the two-break minimum LM test in the present application.

[9]Augmented Dickey-Fuller tests reject the null hypothesis of a unit root at the one-percent level in each of the de-trended series.

[10]The de-trended series $PREM^{det}$, RX^{det}, and DI^{det} are the estimated residuals, $\hat{e}_t$, from the regression

$$y_t = m_0 + m_1 D_t(T_{B1}) + m_2 D_t(T_{B2}) + m_3 t + m_4 DT_t(T_{B1}) + m_5 DT_t(T_{B2}) + e_t$$

for y = *PREM*, *RX*, and *DI*, respectively. Lagged AR error terms are included in the equation to correct for serial correlation, with the chosen lag length determined by the "general to specific" method (described in the Appendix) starting with a maximum of 24 lags. $D_t(T_{Bj})$ and $DT_t(T_{Bj})$, j = 1, 2, are defined in equation (A1) in the Appendix, with the estimated break dates, T_{Bj}, j = 1, 2, taken from Table 2.

$$PREM_t^{det} = b_0 + b_1 RX_t^{det} + b_2 DI_t^{det} + \boldsymbol{b'MONTH} + v_t. \qquad (2)$$

Equation (2) is estimated using OLS with a lagged AR error structure (with the optimal lag length determined using the "general to specific" method described above starting with a maximum of 24 lags) to correct for serial correlation. White's robust standard errors are utilized to correct for possible heteroskedasticity in the error terms.

Column 1 of Table 3 reports the results from the estimation of (2) using OLS. Both the real exchange rate and the interest differential have the expected sign and are statistically significant. A real depreciation (or increase in *RX*) leads to a decline in the black-market premium. An increase in U.S. interest rates relative to those in Yugoslavia, adjusted for official depreciation, leads to an increase in the black-market premium. The equation explains a substantial part of the variation in the premium.

The role of seasonal factors is assessed with the monthly dummies included in the estimation of (2). The constant term shows a premium of 1.46 percent for the base month of January. The estimated coefficients for the other months indicate the amount by which the premium in that month exceeds the average for January. The monthly dummies yield the expected pattern of seasonal variation over the year. The tourist season, which peaks during the summer months, is shown by a seasonal decline of the premium during the months of May through July. The premium during these months is found to be statistically significantly below the January average, with the seasonal peak occurring in June with a decline of nearly 6 percent from the January level. An *F*-test reveals that, as a group, the monthly seasonal dummies are significantly different from zero at the one-percent level.[11]

V. Variability in the Black-Market Premium

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[11] Shachmurove (1999) was unable to find evidence of seasonal variation in a panel of developing countries.

Visual inspection of Figure 1 suggests a structural change in the behavior of the black-market premium in the early 1980s. Indeed, the formal trend-break tests of the previous section indicate a structural break in May of 1982, just two years after the death of Tito and coincident with legal changes increasing government involvement in the black market. The downward trend in the premium following the May, 1982 break (as shown in Figure 1) is consistent with a period of increased government intervention in the black market in the early 1980s. Specifically, the government crack-down on black market activities (which begins with the May, 1982 legal changes) to come up with hard currency in the treasury gives rise to an accumulation of dollars and a decline in the premium.

The black-market premium also exhibits an increase in volatility during the early 1980s. While this aspect of structural change is not captured by the trend-break tests of the previous section, an increase in the volatility of the premium is also consistent with an era of increased government involvement in black-market activities. Recall from Section II the history of black-market activities in Yugoslavia which ends with a period of increased government involvement. Phylaktis (1992) investigates the effects of government-imposed foreign exchange restrictions on the black-market premium in Chile, finding that such restrictions are important determinants of the black-market premium. Since detailed quantitative data on government restrictions are not available for the Yugoslav experience, the following analysis suggests an alternative approach based on trend break evidence combined with knowledge of the timing of regime change and important legal changes affecting black-market activity.

In the remainder of this section, we consider a model of the black market in Yugoslavia which allows an explicit specification of the conditional variance, or volatility, of the black-market premium. Bollerslev (1986) introduced the class of generalized autoregressive conditional heteroskedasticity (GARCH) models, where the variance of the dependent variable is modeled as a function of past values of the dependent variable and independent, or exogenous, variables. In particular, we estimate the following standard GARCH(p,q) specification:

$$PREM_t^{det} = b_0 + b_1 RX_t^{det} + b_2 DI_t^{det} + \boldsymbol{b}'\boldsymbol{MONTH} + v_t, \tag{3}$$

$$s_t^2 = w + a_1 v_{t-1}^2 + a_2 v_{t-2}^2 + ... + a_p v_{t-p}^2 + c_1 s_{t-1}^2 + c_2 s_{t-2}^2 + ... + c_q s_{t-q}^2 + \boldsymbol{r}'\boldsymbol{x}_t, \tag{4}$$

where s_t^2 is the one-period ahead forecast variance based on past information, also known as the conditional variance. We will refer to a_i, i = 1, ..., p, as the ARCH(i) term, and c_j, j = 1, ..., q, as the GARCH(j) term. The vector $\boldsymbol{x}$ is a set of predetermined, or exogenous, regressors in the variance equation, (4). Note that the mean equation, (3), is the same model estimated earlier in equation (2).

The results from estimating the model in (3) and (4) are reported in column 2 of Table 3. A set of monthly dummies is included (in $\boldsymbol{x}$) to assess the influence of seasonal factors in the volatility of the black-market premium. Also in $\boldsymbol{x}$, we include a dummy variable, denoted D(1982:05), which equals one beginning in May of 1982 to test the hypothesis that an increase in volatility was associated with regime change during the early 1980s.[12] The use of the GARCH model leaves the estimated coefficients on the real exchange rate, interest differential, and seasonal factors in the mean equation unaffected. The estimate of the variance equation is obtained using a GARCH(3,3) model, although the estimated ARCH and GARCH coefficients are not reported to conserve space.[13] The estimate of the variance equation suggests an

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[12]May, 1982 is the first estimated break date in the premium series obtained in the analysis of Table 2.

[13] The chosen order (p, q) of the GARCH model reported in Table 3 is determined in a manner similar to the “general to specific” method discussed previously, starting with a GARCH(9,9) model. Lagged AR terms are included in the mean equation, (3), to correct for serial correlation, with the chosen number of lags determined by the “general to specific” method starting with a maximum of 24 lags. The particular model reported in Table 3 is an illustration of the kind of results obtained using such a model. The results are qualitatively similar using other specifications of the general model.

important seasonal effect on the volatility of the black-market premium, as several of the monthly dummies are statistically significant at least at the 10% level. As a group, the monthly seasonal dummies are significant at the one-percent level.

The regime change dummy, *D*(1982:05), is also significant at the one-percent level, and the estimated coefficient suggests an increase in the volatility of the black-market premium beginning in the early 1980s. The estimated conditional standard deviation of the premium is shown in Figure 4. The figure shows the apparent seasonal aspect of the volatility, and also indicates an increase in the mean level of volatility beginning in the early 1980s. More formal evidence on the increase in volatility is obtained by performing a one-break version of the Lee and Strazicich (2003) trend-break test on the estimated conditional standard deviation series depicted in Figure 4 (see Lee and Strazicich (2004) for details). The results indicate a statistically significant break in November, 1982, shortly after the May, 1982 adoption of the new law empowering state activity in the black market (recall the discussion in Section II). The estimated trend of the conditional standard deviation series is shown in Figure 4 along with the original series. Casual inspection suggests an increase of approximately 33% in the mean level of volatility associated with the November, 1982 break.

Such an increase in volatility suggests an important risk perceived by speculators that makes dinars and dollars less perfect substitutes. This behavior during the early 1980s is consistent with an era of increased government involvement in the black market following the death of Tito. Thus, the evidence supports the hypothesis that increased government involvement in the black market increased the volatility of the premium, just as the underlying model suggests.

VI. Conclusion

We apply the Dornbusch *et al* (1983) model of the black-market premium to the black market for dollars in Yugoslavia over the period 1974-1987. We first apply tests of structural change to address the nonstationarity of the variables in the model, an approach not considered in previous applications of the model. Our finding is that all of the variables are trend-break stationary, with estimated break dates closely associated with regime change and legal changes following the death of Tito in the early 1980s. We use this information to de-trend the variables before estimating the model. Once the variables are rendered stationary, we find strong support for the underlying model, including an important seasonal component associated with the summer tourist season in Yugoslavia.

We extend our analysis to explain the increase in volatility that is observed in the black-market premium during the early 1980s. Using a standard GARCH specification, we find evidence of a significant seasonal component in the volatility of the black-market premium. In addition, the evidence suggests a significant increase in the volatility of the premium during the early 1980s. This increase coincides with the beginning of an era of increased government involvement in black-market activities in Yugoslavia, suggesting that government involvement reduced the substitutability between the official and black-market currencies.

Appendix: Review of Formal Tests for Structural Change

Perron (1989) proposed the following methodology for testing the unit root null hypothesis against the trend-break stationary alternative: (1) specify the location of the break date (T_B); (2) estimate a regression that nests the random walk null and the trend-break stationary alternative with either a change in intercept (Model A), a break in slope (Model B), or both (Model C); and (3) use the t-statistic on the first lag of the dependent variable to carry out the test of the null hypothesis. This type of test is often referred to as an "exogenous" break unit root test since the break data must be specified a priori.

Christiano (1992), among others, criticizes the assumption that the location of the break is known a priori in Perron's methodology. Christiano shows that if the break date is not known and the researcher determines the location of the break by visually inspecting the data, the unit root null will be rejected too often. This criticism gave rise to an extension of Perron's methodology which does not require pre-specification of the break date. The strategy applies Perron's methodology for each possible break date in the sample, yielding a sequence of t-statistics. From this sequence, various algorithms can be used to construct 'minimum-t-statistics' which maximize evidence against the null hypothesis. One example is to use the minimum of the sequence of t-statistics as proposed by Zivot and Andrews (1992). As such, the Zivot and Andrews test determines the break point where the unit-root test statistic is the most negative and, therefore, the least favorable to the null hypothesis.[14]

A potential problem with these augmented Dickey-Fuller type endogenous break unit root tests is that they derive their critical values assuming no break under the null hypothesis. Nunes *et al* (1997) and Lee and Strazicich (2001) provide evidence that assuming no break under the null causes the test statistic to diverge and leads to too many rejections of the unit root null when the true data-generating process is a unit root with break(s). To prevent such "spurious rejections," Lee and Strazicich (2003) propose an

[14] Although the location of the break need not be specified a priori, the form of the break under the alternative hypothesis (i.e., Model A, B, or C) must be specified. Sen (2003) shows that a loss of power occurs when the form of the break is incorrectly specified under the alternative, and suggests using the more general Model C when the form of the break is unknown. Thus, we use Model C in our analysis below.

endogenous two-break unit root test. This two-break "minimum LM" unit root test does not diverge in the presence of breaks under the null hypothesis, so that rejection of the null unambiguously implies trend-break stationarity.

We use the methodology of Lee and Strazicich (2003) to test for structural breaks in the series *PREM*, *RX*, and *DI*. It seems particularly appropriate to use a test that includes the possibility of breaks under the null hypothesis given our prior knowledge of regime change (and, thus, the possibility of structural breaks) in the underlying data. The two-break "minimum LM" unit root test statistic can be estimated by regression according to the LM (score) principle as follows:

$$dy_t = l' dZ_t + hS_{t-1} + g_1 dS_{t-1} + g_2 dS_{t-2} + ... + g_k dS_{t-k} + e_t, \qquad \text{(A1)}$$

where S_t is a detrended series such that $S_t = y_t - P^* - Z_t l^*$ for $t = 2, ..., T$; l^* are coefficients in the regression of dy_t on dZ_t with d defined as the first-difference operator; P^* is given by $y_1 - Z_1 l^*$ where y_1 and Z_1 denote the first observations of y_t and Z_t, respectively; and e_t is the contemporaneous error term and is assumed to be independent and identically distributed with zero mean and finite variance. Z_t is a vector of exogenous variables defined by the data-generating process which, corresponding to Model C, is described by $[1, t, D_t(T_{B1}), D_t(T_{B2}), DT_t(T_{B1}), DT_t(T_{B2})]'$, where $D_t(T_{Bj}) = 1$ for $t >= T_{Bj} +1$, $j = 1, 2$, and zero otherwise, $DT_t(T_{Bj}) = t - T_{Bj}$ for $t >= T_{Bj} +1$, $j = 1, 2$, and zero otherwise, and T_{Bj} represents the time period of the breaks. Finally, dS_{t-i}, $i = 1, ..., k$, terms are included as necessary to correct for serial correlation. Note that (A1) involves dZ_t instead of Z_t, so that dZ_t becomes $[1, B_{1t}, B_{2t}, D_{1t}, D_{2t}]'$, where $B_{jt} = dD_t(T_{Bj})$ and $D_{jt} = dDT_t(T_{Bj})$, $j = 1, 2$.

The unit root null hypothesis is described in (A1) by $h = 0$ and the LM test statistic is given by:

$$t^* = t\text{-statistic for the null hypothesis } h = 0.$$

To endogenously determine the location of two breaks ($n_j = T_{Bj}/T$, $j = 1, 2$), the minimum LM unit root test uses a grid search as follows:

$$\mathrm{LM}_{t^*} = \mathrm{Inf}_n t^*(n).$$

Lee and Strazicich (2003) report critical values for Model C, which depend somewhat on the location of the breaks (n_j).

To determine the value of k, the number of dS_{t-i} terms needed to correct for serial correlation, we use the following "general to specific" method. At each combination of break points $n = (n_1, n_2)'$ in the time interval $[0.1T, 0.9T]$ (to eliminate endpoints), we begin with a maximum of 24 lags of dS and examine the statistical significance of the last term. If the last term is not significantly different from zero at the 10% level (using the asymptotic normal distribution), the last lagged term is dropped and the model re-estimated with $k = 23$ lags, and so on until either the last term is significant or $k = 0$. This procedure has been shown to perform well relative to other data-dependent procedures; see Ng and Perron (1995). As such, we jointly determine the location of breaks and the number of lagged dS terms endogenously from the data.[15]

[15]Gauss code for the one- and two-break minimum LM unit root test is available at http://www.cba.ua.edu/~jlee/gauss.

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Table 1
Mean and Standard Deviation of Model Variables

Variable	1974:1-1987:12	1974:1-1982:9	1982:10-1987:12
PREM			
Mean	10.22	7.52	14.60
Std. Dev.	10.10	6.59	12.96
RX			
Mean	30.56	22.90	43.01
Std. Dev.	11.01	3.18	7.07
DI			
Mean	-10.21	5.05	-34.77
Std. Dev.	27.76	6.31	31.19

Table 2
Two-Break Minimum LM Unit Root Tests

Series	k	T_{B1}, T_{B2}	Test Statistic	$n = (T_{B1}/T, T_{B2}/T)$
PREM	0	1982:05, 1986:06	-7.98***	(0.6, 0.8)
RX	23	1980:05, 1984:01	-5.72**	(0.4, 0.8)
DI	22	1983:10, 1985:07	-9.16***	(0.6, 0.8)

Notes: k is the optimal number of lagged first-difference terms included in the unit root test to correct for serial correlation. T_{B1} and T_{B2} are the estimated break dates. n is the location of the breaks within the sample for which critical values are determined. Critical values are reported in Table 1 of Strazicich *et al* (2004). ***, **, and * indicate significance at the 1%, 5%, and 10% levels, respectively.

Table 3
Yugoslav Black Market Premium Model, 1974-1987
Dependent Variable: $PREM^{det}$

	Estimation by:	
Variable	OLS	GARCH(3,3)
Constant	1.46	1.14
RX^{det}	-2.03***	-1.93***
DI^{det}	0.54***	0.49***
Feb	2.71**	2.73***
Mar	-0.63	-1.28
Apr	-1.01	-1.37
May	-2.84*	-2.98***
Jun	-5.79***	-5.27***
Jul	-4.31**	-3.84***
Aug	-1.37	-1.46
Sep	-0.39	-0.94
Oct	1.21	0.47
Nov	-1.79	-1.40
Dec	-2.22	-1.87
Lagged AR terms	1	21
Adjusted R^2	0.565	0.379
Durbin-Watson	2.002	1.880
F-test	3.537***	12.171***
	Variance Equation	
Constant	--	13.23**
Feb	--	-13.10
Mar	--	-0.84
Apr	--	-14.73**
May	--	-19.29***
Jun	--	1.83
Jul	--	-11.02
Aug	--	12.75**
Sep	--	-13.59**
Oct	--	-23.73***
Nov	--	-5.71
Dec	--	-10.23
D(1982:05)	--	7.38***
F-test	--	3.230***
ARCH LM test	--	0.43

Notes: See variable definitions in the text. *, **, and *** denote significance at the 10%, 5%, and 1% levels, respectively, using Bollerslev-Wooldrige robust standard errors. *F*-test is the *F*-statistic for a test of the restriction that all monthly coefficients are equal to zero in each panel of the table. ARCH LM test is the *F*-statistic for a test of the null hypothesis that there is no additional ARCH up to 24 lags in the residuals.

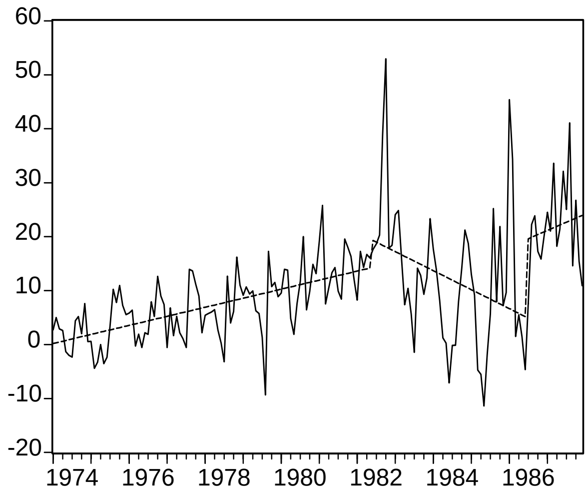

Figure 1: Black Market Premium (%)

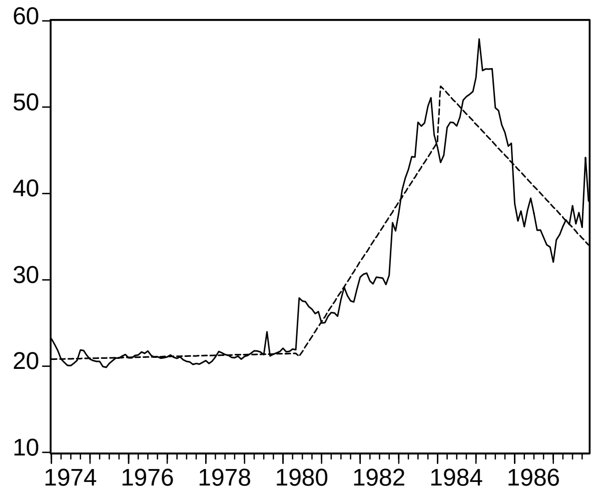

Figure 2: Real Exchange Rate (dinar/dollar)

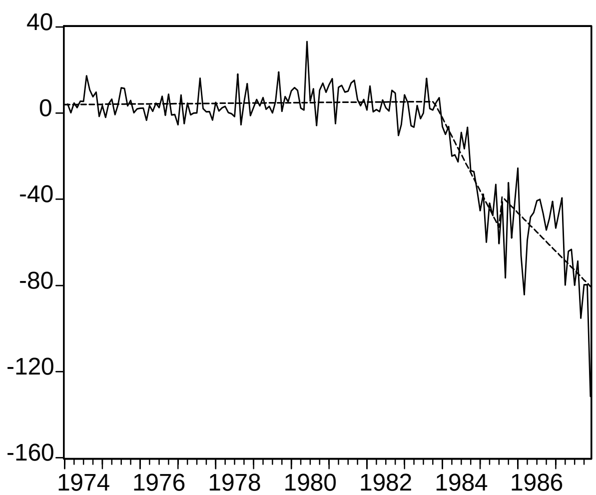

Figure 3: Depreciation-adjusted Interest Differential

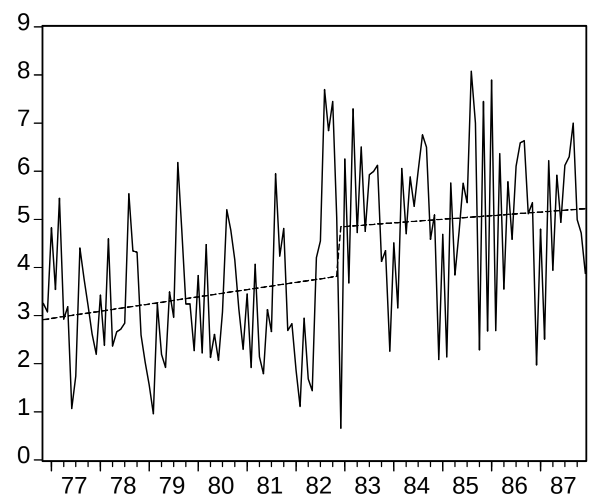

Figure 4: Estimated Conditional Standard Deviation of the Black Market Premium