Trend breaks and non-stationarity in the Yugoslav black market for dollars, 1974–1987

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

ABSTRACT

We estimate a model of the black market premium for dollars in Yugoslavia from 1974 to 1987. Unlike previous applications of the model, our analysis addresses non-stationarity 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.
I. INTRODUCTION

This paper applies the model of Dornbusch et al. (1983) to the black market for dollars in Yugoslavia from 1974 to 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 non-stationarity 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 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. The 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 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 1982, a new law empowered the state with 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 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 cheques 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.

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

The model

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

\[ PREM_t = b_0 + b_1 RX_t + b_2 DI_t + b'MONTH + u_t, \] (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, \( MONTH=[FEB, \ldots, DEC] \) is a vector of monthly dummy variables (January
is omitted because a constant is included),
b=[b3, . . . , b13], 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 b1<0
and b2>0 in Equation 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. 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.

The data

Monthly data on Yugoslavia’s black market
exchange rate are taken from World Currency
Yearbook (International Currency Analysis, Inc.,
Various). Other data are from the International
Financial Statistics. 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 US producer price index to the Yugoslav
producer price index to calculate the official real
exchange rate, RX. Short-term (three-month) nominal
interest rates are used to calculate the depreciation-
adjusted interest differential, DI, which is
defined as iUSpd_iYugo, where iUS and iYugo 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 to 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 unit root 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 non-stationary, 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. Table 1 also suggests a change in behaviour in the early 1980s, as the mean and standard deviation of each series is noticeably different before and after this period.

**Evidence on trend breaks in the data**

As noted above, conventional Dickey–Fuller tests imply that each of the series shown in Figs 1–3 are non-stationary 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 non-stationary using conventional unit root tests.
Fig. 1. Black market premium (%)

Fig. 2. Real exchange rate (dinar/dollar)
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). The results are summarized in Table 2. For all three series, the null hypothesis of a unit root is rejected in favour of the trend break stationary alternative. Thus, in using these series to estimate the model in Equation 1, de-trending is appropriate to render the series stationary. The estimated trends for the PREM, RX and DI series are shown in Figs. 1, 2 and 3, respectively, along with the original series.

Table 1. Mean and standard deviation of model variables

<table>
<thead>
<tr>
<th></th>
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</thead>
<tbody>
<tr>
<td><strong>PREM</strong></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Mean</td>
<td>10.22</td>
<td>7.52</td>
<td>14.60</td>
</tr>
<tr>
<td>Std. dev.</td>
<td>10.10</td>
<td>6.59</td>
<td>12.96</td>
</tr>
<tr>
<td><strong>RX</strong></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Mean</td>
<td>30.56</td>
<td>22.90</td>
<td>43.01</td>
</tr>
<tr>
<td>Std. dev.</td>
<td>11.01</td>
<td>3.18</td>
<td>7.07</td>
</tr>
<tr>
<td><strong>DI</strong></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Mean</td>
<td>-10.21</td>
<td>5.05</td>
<td>-34.77</td>
</tr>
<tr>
<td>Std. dev.</td>
<td>27.76</td>
<td>6.31</td>
<td>31.19</td>
</tr>
</tbody>
</table>

Fig. 3. Depreciation-adjusted interest differential
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.

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 non-stationary. 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. We denote the resulting de-trended stationary series as PREM$_{det}$, RX$_{det}$ and DI$_{det}$, respectively. Thus, the model to be estimated is:

$$PREM_t^{det} = b_0 + b_1 RX_t^{det} + b_2 DI_t^{det} + b'MONTH + v_t.$$  \hspace{1in}(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 Equation 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 US 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.
Table 3. Yugoslav black market premium model, 1974–1987

<table>
<thead>
<tr>
<th>Variable</th>
<th>Estimation by OLS</th>
<th>Estimation by GARCH(3, 3)</th>
</tr>
</thead>
<tbody>
<tr>
<td>Constant</td>
<td>1.46</td>
<td>1.14</td>
</tr>
<tr>
<td>RX_d$t$</td>
<td>$-2.03^{**}$</td>
<td>$-1.93^{**}$</td>
</tr>
<tr>
<td>Dr_d$t$</td>
<td>$0.54^{***}$</td>
<td>$0.49^{***}$</td>
</tr>
<tr>
<td>Feb</td>
<td>$2.71^{**}$</td>
<td>$2.73^{***}$</td>
</tr>
<tr>
<td>Mar</td>
<td>$-0.63$</td>
<td>$-1.28$</td>
</tr>
<tr>
<td>Apr</td>
<td>$-1.01$</td>
<td>$-1.37$</td>
</tr>
<tr>
<td>May</td>
<td>$-2.84^{*}$</td>
<td>$-2.98^{***}$</td>
</tr>
<tr>
<td>Jun</td>
<td>$-5.79^{***}$</td>
<td>$-5.27^{***}$</td>
</tr>
<tr>
<td>Jul</td>
<td>$-4.31^{**}$</td>
<td>$-3.84^{***}$</td>
</tr>
<tr>
<td>Aug</td>
<td>$-1.37$</td>
<td>$-1.46$</td>
</tr>
<tr>
<td>Sep</td>
<td>$-0.39$</td>
<td>$-0.94$</td>
</tr>
<tr>
<td>Oct</td>
<td>$1.21$</td>
<td>$0.47$</td>
</tr>
<tr>
<td>Nov</td>
<td>$-1.79$</td>
<td>$-1.40$</td>
</tr>
<tr>
<td>Dec</td>
<td>$-2.22$</td>
<td>$-1.87$</td>
</tr>
<tr>
<td>Lagged AR terms</td>
<td>1</td>
<td>21</td>
</tr>
<tr>
<td>Adjusted $R^2$</td>
<td>0.565</td>
<td>0.379</td>
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<tr>
<td>Durbin–Watson</td>
<td>2.002</td>
<td>1.880</td>
</tr>
<tr>
<td>$F$-test</td>
<td>3.537***</td>
<td>12.171***</td>
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</tbody>
</table>

Variance equation

<table>
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<tr>
<th>Variable</th>
<th>OLS</th>
<th>GARCH(3, 3)</th>
</tr>
</thead>
<tbody>
<tr>
<td>Constant</td>
<td>$-13.23^{**}$</td>
<td>$-13.10$</td>
</tr>
<tr>
<td>Feb</td>
<td>$-13.10$</td>
<td>$-0.84$</td>
</tr>
<tr>
<td>Mar</td>
<td>$-14.73^{**}$</td>
<td>$-19.29^{***}$</td>
</tr>
<tr>
<td>Apr</td>
<td>$-19.29^{***}$</td>
<td>$1.83$</td>
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<tr>
<td>May</td>
<td></td>
<td>$-11.02$</td>
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<td>Jun</td>
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<td>$12.75^{**}$</td>
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<td>Jul</td>
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<td>$-13.59^{**}$</td>
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<td>Aug</td>
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<td>Oct</td>
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<tr>
<td>Nov</td>
<td></td>
<td>$7.38^{***}$</td>
</tr>
<tr>
<td>Dec</td>
<td></td>
<td>$3.230^{***}$</td>
</tr>
<tr>
<td>Dt(1982:05)</td>
<td></td>
<td>$0.43$</td>
</tr>
</tbody>
</table>

Notes: Dependent variable: $PREM_d$t See variable definitions in the text.
* ** and *** denote significance at the 10%, 5% and 1% levels, respectively, using Bollerslev–Wooldridge 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.
The role of seasonal factors is assessed with the monthly dummies included in the estimation of Equation 2. The constant term shows a premium of 1.46% 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% from the January level. An F-test reveals that, as a group, the monthly seasonal dummies are significantly different from zero at the 1% level.11

V. VARIABILITY IN THE BLACK MARKET PREMIUM

Visual inspection of Fig. 1 suggests a structural change in the behaviour 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 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 Fig. 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 modelled 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:

\[
PREM_t^{det} = b_0 + b_1 RX_t^{det} + b_2 DR_t^{det} \\
+ b'MONTH + v_t, \tag{3}
\]

\[
s_t^2 = w + a_1 v_{t-1}^2 + a_2 v_{t-2}^2 + \cdots + a_p v_{t-p}^2 \\
+ c_1 s_{t-1}^2 + c_2 s_{t-2}^2 + \cdots + c_q s_{t-q}^2 + r'x_t, \tag{4}
\]

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 \(a_i, i=1, \ldots, p,\)
as the ARCH\((i)\) term, and \(c_j, j=1, \ldots, q,\) as the
GARCH\((j)\) term. The vector \(x\) 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 Equations
3 and 4 are reported in column 2 of Table 3. A set
of monthly dummies is included \((in x)\) to assess
the influence of seasonal factors in the volatility of
the black market premium. Also in \(x\), we include a
dummy variable, denoted \(D(1982:05)\), which equals
one beginning in May 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. The estimate of the variance equation suggests an 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 1% level.

The regime change dummy, \( D(1982:05) \), is also significant at the 1% 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 Fig. 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 Fig. 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 Fig. 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 behaviour 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.
ACKNOWLEDGEMENT

We thank Amit Sen for useful comments on an earlier draft.

NOTES

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 non-stationarity. Other studies use a cointegration approach to address non-stationarity 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.

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

4. An increase in RX denotes a real depreciation.

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.


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 1% level in each of the de-trended series.

10. The de-trended series $\hat{PREM}_{ret}$, $\hat{RX}_{ret}$ and $\hat{DI}_{ret}$ are the estimated residuals, $\hat{\epsilon}_t$, from the regression $\hat{\epsilon}_t = m_0 + m_1 D_{t}(T_B) + m_2 D_{t}(T_B) + m_3 D_{t}(T_B) + m_4 D_{t}(T_B) + \epsilon_t$ 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. $D_{t}(T_B)$ and $D_{t}(T_B)$, $j = 1, 2$, are defined in Equation (A.1) in the Appendix, with the estimated break dates, $T_B$, $j = 1, 2$, taken from Table 2.

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

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

REFERENCES


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 date 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 prespecification 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 favourable 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 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:
\[ \Delta y_t = \delta' \Delta Z_t + \phi S_{t-1} + \sum \gamma_i \Delta S_{t-i} + \epsilon_t \quad (A1) \]

where \( S_t \) is a detrended series such that \( S_t = y_t \Psi - Z_t \delta \) for \( t = 2, \ldots, T \), \( \delta \) are coefficients in the regression of \( \Delta y_t \) on \( \Delta Z_t \) with \( \Delta \) defined as the first-difference operator; \( \Psi \) is given by \( y_1 - Z_1 \delta \) where \( y_1 \) and \( Z_1 \) denote the first observations of \( y_t \) and \( Z_t \) respectively; and \( \epsilon_t \) 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 \([1, t, D_t(T_{B1}), D_t(T_{B2}), DT_t(T_{B1}), DT_t(T_{B2})]^T\), where \( D_t(T_{Bj}) = 1 \) for \( t \geq T_{Bj} + 1, j = 1, 2, \) and zero otherwise, \( DT_t(T_{Bj}) = t - T_{Bj} \) for \( t \geq T_{Bj} + 1, j = 1, 2, \) and zero otherwise, and \( T_{Bj} \) represents the time period of the breaks. Finally, \( \Delta S_{t-i}, i = 1, \ldots, k, \) terms are included as necessary to correct for serial correlation. Note that Equation A1 involves \( \Delta Z_t \) instead of \( Z_t \), so that \( \Delta Z_t \) becomes \([1, B_{1t}, B_{2t}, D_{1t}, D_{2t}]^T\), where \( B_{jt} = \Delta D_t(T_{Bj}) \) and \( D_{jt} = \Delta DT_t(T_{Bj}), j = 1, 2. \)

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

\[ \tau = t\text{-statistic for the null hypothesis } \phi = 0. \]

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

\[ LM_\tau = \text{Inf}_\lambda \tau(\lambda). \]

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

To determine the value of \( k \), the number of \( \Delta S_{t-i} \) terms needed to correct for serial correlation, we use the following 'general to specific' method.

At each combination of break points \( \lambda = (\lambda_1, \lambda 2) \)' in the time interval \([0.1T, 0.9T] \) (to eliminate endpoints), we begin with a maximum of 24 lags of \( \Delta S \) 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 $\Delta S$ terms endogenously from the data.