Private sector credit and inflation during Brazilian stabilization plans: Models with endogenously determined structural breaks

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Fabiana Rocha

Abstract

The purpose of this paper is to investigate the relationship between credit, inflation, and stabilization plans in Brazil during 1970-1998. The structural changes suffered by credit are endogenously determined using Bai and Perron's methodology (1998). The breaks correspond to the major stabilization plans adopted by the country. A negative relationship between credit and inflation is found. Stability contributes to credit expansion without, however, guaranteeing its quality.

Resumo

O objetivo deste artigo é investigar a relação entre crédito, inflação e planos de estabilização no Brasil durante o período 1970-1998. Determina-se endogêneamente as mudanças estruturais sofridas pelo crédito, usando-se a metodologia de Bai e Perron (1998). As quebras coincidem com os principais planos de estabilização adotados pelo país. Acha-se uma relação negativa entre crédito e inflação, indicando que a estabilidade contribui para a expansão do crédito embora não garanta a sua qualidade.

Key Words: Credit, inflation, stabilization plans, structural breaks.

JEL Code: E5, C22.

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1. Introduction.

The literature on stabilization plans has given little attention to credit behavior. In general, the studies focus on the behavior of consumption and product, and on the dynamics of convergence for interest rates\(^1\). De Gregório, Guidotti and Végh (1998) seem to be an exception, since they are the only ones who refer, albeit superficially, to credit. They have shown that the expansion-contraction cycle after the implementation of stabilization plans is particularly evident in the behavior of durable goods. The initial decrease of inflation generates a wealth effect that leads consumers to buy durable goods. As consumers replace their stocks, this effect diminishes. Among the reasons for the occurrence of this cycle, the authors have mentioned the increase in credit supply. From the empirical point of view, Khamis is an exception (1996). He has clearly associated stabilization plans with credit and has concluded that there is a negative relationship between credit and inflation in Mexico, Argentina and Chile, but not in Israel.

The purpose of this paper is also to analyze credit behavior, but taking into account only the Brazilian experience. Besides, contrarily to Khamis (1996), who exogenously determines the dummy variables that represent the stabilization plans, we use Bai and Perron’s (1998a) methodology, in which multiple structural changes are determined endogenously. Subsequently, we evaluate the effect of inflation on credit during each period of structural change.

The article is organized as follows. Section 2 briefly discusses the possible mechanisms whereby inflation reduction contributes to an

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\(^1\) See, for example, Calvo (1986) and Helpman and Razin (1987).
increase in credit supply. Section 3 presents the results of the unit root tests for the inflation and the credit series. Section 4 describes Bai and Perron’s methodology that will be used to study the relationship between credit and inflation. Section 5 presents and discusses the results of the analysis between credit and inflation, whereas section 6 evaluates the gains from including real interest rates in the model. Finally, section 7 summarizes the main conclusions.

2. Credit, stabilization plans and soundness of the banking system.

During the implementation of stabilization plans an increase in the supply of credit and/or in the demand for credit can be observed. A fall in nominal interest rates, although it does not necessarily represent a reduction in real interest rates, when combined with the increase of maturity dates, increases the demand for credit. On the other hand, the end of inflationary gains\(^2\), the reduction of uncertainties about the value of future nominal interest rates in a stable environment and the increase of bank liquidity due to the reduction of nominal interest rates stimulate the expansion of the credit supply.

Table 1 shows the mean and variance of nominal interest rates and inflation rates during the Brazilian stabilization plans.

\(^2\)See Mendonça de Barros and Almeida Junior (1997) and Mendonça de Barros, Loyola and Bogdanski (1998).
Private sector credit and inflation during Brazilian stabilization plans

Table 1
Mean and variance of inflation and interest rates during and between the stabilization plans

<table>
<thead>
<tr>
<th>Period(month/year)</th>
<th>Inflation rates(%)</th>
<th>Nominal interest rates(%)</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>Mean</td>
<td>Variance</td>
</tr>
<tr>
<td>1/1984-2/1986</td>
<td>9.89</td>
<td>3.48</td>
</tr>
<tr>
<td>Cruzado 3/1986-10/1986</td>
<td>1.81</td>
<td>0.47</td>
</tr>
<tr>
<td>Bresser 7/1987-9/1987</td>
<td>6.81</td>
<td>5.74</td>
</tr>
<tr>
<td>10/1987-1/1989</td>
<td>20.28</td>
<td>37.24</td>
</tr>
<tr>
<td>5/1989-2/1990</td>
<td>44.78</td>
<td>454.22</td>
</tr>
<tr>
<td>6/1990-1/1989</td>
<td>16.03</td>
<td>14.29</td>
</tr>
<tr>
<td>Collor II 2/1991-6/1991</td>
<td>7.55</td>
<td>2.77</td>
</tr>
<tr>
<td>Real 7/1994-12/1998</td>
<td>0.94</td>
<td>1.74</td>
</tr>
</tbody>
</table>

Note: The duration of each plan follows the dates adopted by Cati, Garcia and Perron (1999).

Source: Selic interest rates and IPC-Fipe monthly inflation rates.

In periods of stabilization or decreases in inflation (lower mean), the variances of inflation and nominal interest rates are lower than in periods of high inflation. As observed by Münch (1998), the Brazilian inflation rate explains part of the volatility of the nominal interest rate, and the latter one has a negative impact on credit.

The high default rate after the stabilization plans, despite the improvement on the evaluation of risks implies that part of the credit...
expansion was performed on bad loans. This happened for two reasons. Firstly, a moral hazard problem, when people become less cautious in managing their debts given that they believe they will not bear alone the total cost of their choice. Secondly, an adverse selection problem. Due to the increase in credit, the banks end up lending to clients with high-risk projects. Since bad credit expansion can weaken the financial system and, therefore, the economy as a whole, the creation of regulatory and prudential policies would have been necessary in order to avoid the bank insolvency problems that followed, especially after the Real Plan.

3. Order of integration of inflation and credit.

In this section we describe the data used and we determine the order of integration of the Brazilian inflation and private sector credit series.

3.1. Order of integration of inflation.

The adoption of several and successive stabilization plans in Brazil makes the task of testing for the presence of a unit root in the Brazilian inflation series nontrivial. Cati, Garcia and Perron (1999) show that in the presence of inliers, the conventional unit root tests tend to reject the hypothesis of non-stationarity. By using the monthly series of Brazilian inflation rates from January 1974 to June 1993 and by considering their inliers, these authors show that inflation has a unit root in this period. However, the series used

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3Lindgren, Garcia and Saal (1996), by studying eight countries that have been through bank crises- Argentina, Chile, Finland, Ghana, Norway, Philippines, Uruguay and Venezuela, observed that most times there was an excessive credit expansion to the private sector before the crises.

4There is an inlier when the series has a swift downward movement before moving back to the previous trend line.
herein starts on January 1970 and ends on December 1998 (Graph 1). A question that arises is whether the inclusion of these additional years would change the behavior of the inflation series.

Graph 1: monthly inflation rate

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5 Cati, Garcia and Perron (1999) built a monthly inflation series from different indices considered official over time. Given that the purpose of this study is to analyze the relationship between credit and inflation during the stabilization plans, we would rather use the IPC-Fipe, once it has never been purged in any of the inflation control programs adopted. We should observe that, since the series has no seasonality, the same results obtained for the monthly series are valid for the quarterly data.
Intuitively we presumed that the series from 1970 to 1998 would remain integrated of order 1 (I(1)) for the following reasons:


2) Concerning the period after the Real Plan (1994-1998), we would be adding an additive outlier, that is, a change in the intercept of the series. If we removed the effect of this structural break we would have an extension of Cati, Garcia and Perron (1999)'s series with no breaks.

In order to confirm the intuition above, we used the same Cati, Garcia and Perron (1999)'s ADF(CA) test, and we concluded that the monthly inflation series from January 1970 to December 1998 is I(1) (Table 2).

<table>
<thead>
<tr>
<th>Table 2</th>
</tr>
</thead>
<tbody>
<tr>
<td><strong>ADF(CA) test for inflation</strong></td>
</tr>
<tr>
<td>ADF(CA)</td>
</tr>
<tr>
<td>( k ) determined by the BIC criteria as suggested by Cati, Garcia and Perron (1999)</td>
</tr>
</tbody>
</table>

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6The ADF test modified by Cati, Garcia and Perron (1999) adds dummy variables related to the beginning, duration and end of the stabilization plans (see dates in Table 1) to the traditional ADF test, i.e., \( y_t = \mu + \beta t + \sum (dummies - beginning, duration and end of stabilization plans) + \alpha y_{t-1} + \sum_{j=1}^{K} c_j \Delta y_{t-j} + v_t \). The ADF(CA) statistics is built as the \( t \) statistics to test whether the sum of the autoregressive coefficients (\( \alpha \)) is equal to one.

7The ADF(CA) test does not reject the presence of unit root for the level of the inflation series (critical values are the same as those of traditional ADF test), rejects the presence of unit root for the inflation series in first difference at 1%.
3.2. Order of integration of credit.

We considered the credit series as the real amount of credit granted to the private sector of the total available credit deposited in the banks, divided by the Gross Domestic Product (GDP). We used quarterly data from the International Monetary Fund – *International Financial Statistics* for the period starting in the first quarter of 1970 and ending in the last quarter of 1998 (1970:1-1998:4). The series includes consumer credit and credit for investments. However, the percentage corresponding to each one of them is unknown.

Graph 2 shows the Brazilian private sector credit as a ratio of GDP (from now on we will refer to it simply as “credit”) as well as the different stabilization plans adopted in the country from 1970 up to 1998.
The series remains at about 0.10 until the first semester of 1979 when it drops to an average of 0.08 until 1985. Therefore, in 1979, we observe a change in the intercept of the trend function. Between 1986 and 1993, a period in which Brazil unsuccessfully implements 5 stabilization plans (Cruzaço, Bresser, Summer, Collor I and Collor II), the series becomes quite volatile. In general, the series reacts positively to each plan, but it returns to the levels observed before the plan only a few quarters later. Finally, after 1994, the trend function reaches a new and higher intercept where it remains until the end of the sample period.

Usually, in order to test for the presence of unit roots ADF and Phillips-Perron tests are performed. However, such tests are not robust to the presence of structural changes in the deterministic trend function. Changes in the level and/or slope of the deterministic trend function, as occurs with the credit series (graph 2), affect the conventional tests, causing them not to reject the unit root hypothesis when, in fact, the series is stationary around a deterministic trend. Therefore, we decided to adopt a test that would allow changes in the trend function. More specifically, as the changes in the deterministic trend function of credit series occur abruptly, we used the additive outlier model (AO)\(^8\) proposed by Perron (1994a) and Vogelsang and Perron (1994). We applied it to the full sample (1970-1998) and to two subsamples to assess the influence of the period prior to the major economic stabilization plans (1970-1985) and the pre-Real Plan period (1970-1994) on the results. Given the characteristics of

\(^{8}\) An additive outlier model is indicated when the changes in deterministic trend function are abrupt while an innovational outlier model should be used in the cases in which changes in the deterministic trend function are gradual.
the series, we used a model with a change in the intercept for the (1970-1985) period and a model with a change in the intercept and in the slope for the other two. In the AO model, the test is performed using a two-step regression procedure. The unit root is tested based on the $t$-statistic for testing that $\alpha = 1$ in the regression:

$$y_t = \alpha y_{t-1} + \sum_{j=1}^{K} c_j \Delta y_{t-j} + e_t; \quad \text{(1)}$$

where $y_t$ represents the original series ($y_t$) without trend, that is, $y_t$ is the residual of the following regression:

a) in the model with one change in the intercept:

$$y_t = \mu + \beta t + \lambda D U_t + y_t; \quad \text{(2)}$$

where $D U_t = 1$ if $t \geq T_b$ and 0 otherwise.

b) in the model with a change in the intercept and in the slope:

$$y_t = \mu + \beta t + \lambda D U_t + \gamma D T^*_t + y_t; \quad \text{where } D T^*_t = t - T_b \text{ if } t \geq T_b \text{ and } 0 \text{ otherwise.} \quad \text{(3)}$$

The break date $T_b$ is determined according to Perron (1994a) and Volgelsang and Perron (1994) and the truncation parameter $k$ is determined following Ng and Perron (1995). To determine $T_b$, it is possible to use two criteria: criterion $\alpha$, to minimize the $t$ statistics associated with $\alpha^* (H_0 : \alpha = 1)$ in equation 1 and criterion $\gamma$,
to maximize or minimize, depending on the direction of the trend function, the t statistics associated with $\gamma^*$ in equations 2 and 3. We used the first criterion.

For the truncation parameter $K$, we used the data dependent rules called $t$-sig and $F$-sig. The procedure begins with a $k$ maximum value (we used $k_{max} = 10$) and with a value of $T_b$ obtained previously.

In the $t$-sig procedure we used the t statistic to test the significance of the last lag of $\Delta y_t$ in equation 1. If the coefficient of $\Delta y_{t-kmax}$ is not significant, the order is reduced until we have a significant coefficient of $\Delta y_{t-j}$ or until we reach the minimum limit $(k = 0)$.

The procedure is similar for the $F$-sig procedure, but it is based on the joint significance of the lag coefficients of the first differences in the series. Two regressions are estimated for equation 1: one with $k_{max}$ and another with $(k_{max}-1)$. Then, we verify whether the $k_{max}$ lag is significant by using the $F$ statistics. If it is not significant, the test is carried out, comparing $(k_{max}-1)$ with $(k_{max}-2)$ and $k_{max}$ with $(k_{max}-2)$ and is repeated until the additional lag is significant.

Using $\alpha$ criterion, the values of the $t$-sig and $F$-sig statistics rejected the unit root hypothesis in all samples, indicating that the credit is stationary around a deterministic trend (Table 3)\(^9\).

\(^9\)Complete table with the test results is found in Zerbini (2000).
Table 3

Unit root tests with structural breaks for credit series

<table>
<thead>
<tr>
<th>Period</th>
<th>AO model with one change in intercept.</th>
<th>$T_b$</th>
<th>$k_{max}$</th>
<th>$K$</th>
<th>$t\alpha$</th>
</tr>
</thead>
<tbody>
<tr>
<td>1970-1985</td>
<td>$t$-sig and $F$-sig</td>
<td>1978:02</td>
<td>10</td>
<td>10</td>
<td>-4.967++</td>
</tr>
<tr>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td></td>
<td>$F$-sig</td>
<td>1986:02</td>
<td>10</td>
<td>10</td>
<td>-3.939</td>
</tr>
<tr>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td></td>
<td>$F$-sig</td>
<td>1985:02</td>
<td>10</td>
<td>9</td>
<td>-4.784+</td>
</tr>
</tbody>
</table>

Notes: ++ and + mean that the unit root hypothesis ($H_0$) is rejected at 5% and 10% respectively.

For the 1970-1985 period, a structural break in the intercept of the deterministic trend function was considered. The test indicated a structural change in the second quarter of 1978 for the $F$-sig and $t$-sig statistics.

For the 1970-1994 and 1970-1998 periods, an additive outlier in the intercept and slope of the deterministic trend function was considered. The test indicated a structural change between the second quarter of 1985 and the fourth quarter of 1987, period in which the Cruzado Plan was implemented. Therefore, after this date, we observe a change in the mean and trend line of the series.

We concluded that, despite all shocks to the credit, its stochastic components remained stationary. The shocks caused by the oil crisis in 1979 and by the adoption of the stabilization plans after 1986 abruptly changed the deterministic trend function, but did not bring...
about a permanent effect. This way, credit expansion policies can be adopted since their effects seem to have limited duration and can, therefore, be controlled.


After studying the order of integration of the credit series, our purpose is to endogenously determine its structural breaks. The literature on structural changes has focused on one break only. The development of models with more than one structural change is relatively recent, having started in 1996. The study conducted by Bai and Perron (1998a) is the first one to propose, albeit at the theoretical level, more flexible hypotheses about the model to be considered. Empirical applications only appear in Bai and Perron (1998b).

The model and the algorithm\textsuperscript{10} proposed by Bai and Perron (1998b) for endogenously determining multiple structural changes start by considering the following multiple linear regression with \( m \) structural changes, or \((m + 1)\) regimes:

\[
\begin{align*}
  y_t &= x_t'\beta + z_t'\delta_1 + \mu_t & t \geq T_1 \\
  y_t &= x_t'\beta + z_t'\delta_2 + \mu_t & T_1 < t \leq T_2 \\
  &\vdots & \vdots \\
  y_t &= x_t'\beta + z_t'\delta_m + \mu_t & T_m < t
\end{align*}
\]

In this model, \( y_t \) is the observed dependent variable at time \( t \), \( x_t(p \times 1) \) and \( z_t(q \times 1) \) are the vector of the explanatory variables, \( \beta \) and \( \delta_j \) (\( j = 1, 2, \ldots, m + 1 \)) are the coefficient vectors, \( \mu_t \) is the random term. The indices \((T_1, T_2, \ldots, T_m)\), or structural breaks, are explicitly treated as unknown.

\textsuperscript{10}Herein we used the algorithm programmed in Gauss, version 2.4, November 19th, 1999.
Private sector credit and inflation during Brazilian stabilization plans

The objective is to estimate the unknown regression coefficients \((\beta \text{ and } \delta_j)\) and the structural break points \(\{T_j\}\) from the known \(T\) observations of \((y_t, x_t, z_t)\).

It is important to observe that if \(p\) is not zero, we have a partial structural change model, the vector \(\beta\) is not subject to changes and is estimated using the full sample. When we have \(p = 0\), we are considering a pure structural change model, in which all coefficients are subject to changes.

The estimation method consists of three parts. First, for each vector with size \((T_1, T_2, \ldots, T_m)\), or \(m\)-partition \(\{T_j\}\), the least squares estimates for vectors \(\beta\) and \(\delta_j\) are obtained by minimizing the sum of the squared residuals:

\[
\sum_{i=1}^{m+1} \sum_{t=T_{i-1}+1}^{T_j} [y_t - x_t'\beta - z_t'\delta_m]^2
\]

Let \(\beta^*(\{T_j\})\) and \(\delta^*(\{T_j\})\) denote the estimates obtained from partition \(\{T_j\}\).

The second step is to determine the optimal partition \(m\{T^*_j\}\). Let the sum of the squared residuals be \(S_T(T_1, T_2, \ldots, T_m)\) and substitute \(\beta^*(\{T_j\})\) and \(\delta^*(\{T_j\})\) in the objective function, the vector of optimal structural break points \(\{T^*_j\}\) is such that

\[
\{T^*_j\} = \{T^*_{i_1}, T^*_{i_2}, \ldots, T^*_{i_m}\} = \arg\min_{\{T_j\}} S_T(T_1, T_2, \ldots, T_m)
\]

where minimization is taken over all partitions \(\{T_j\}\) such that \(T_i - T_{i-1} \geq q\).

Therefore, the structural break point estimators are global minimizers of the objective function. In the last step, given the optimal
\(m\)-partition \(\{T_j^*\}\), the regression parameter estimates are obtained, i.e., \(\beta^*(\{T_j^*\})\) and \(\delta^*(\{T_j^*\})\)\(^{11}\).

Bai and Perron point out that when no lagged dependent variables are allowed in \(\{x_t, z_t\}\), substantial correlation and heteroskedasticity is permitted in the errors \(\{m_t\}\). Otherwise, no serial correlation is allowed. However, in both cases it is possible to have different distributions for the regressors and errors in each segment \([T_i - T_{i-1}]\).

In order to obtain global minimizers of the sum of squared residuals Bai and Perron (1998b) present an algorithm based on the principle of dynamic programming. Their approach emphasizes the fact that, in a sample of size \(T\), this is a problem of order \(O(T^2)\) and not of order \(T^m\), as a standard search algorithm would require for the case of \(m\) breaks.

The order \(O(T^2)\) is due to the fact that the total number of possible segments (length between consecutive breaks) is at most \(T(T + 1)/2\). To understand that, consider a square matrix \(A_T\) in which each element \(a_{ij}\) represents the segment that starts on date \(i\) and ends on date \(j\). Thus, all the possible partitions that start on a break date \(i\) will be represented by elements \(a_{ij}\) with \(j > i\)\(^{12}\). This means that the possible partitions with 1, 2, 3 up to \(T\) breaks will be in the upper triangle of matrix \(A_T\), which has \(T(T+1)/2\) elements. In practice, there are less than \(T(T + 1)/2\) possible segments by simply imposing a minimal length \((h)\) between the breaks or restricting the maximum number \((M)\) of breaks in the series.

The global sum of squared residuals for any \(m\)-partition \(\{T_j\}\)

\(^{11}\)Bai and Perron (1998b) have observed that the \(T_j^*/T\) ratio converges towards its real value \(T_j^*/T\) under a set of hypotheses that are sufficiently general, in such a way that the convergence occurs for a wide variety of models.

\(^{12}\)Partitions with first break on date 3, for example, can only have their second break on date 4, 5, up to \(M\).
and for any value of $m$ must necessarily be a particular linear combination of these $T(T+1)/2$ sums of squared residuals. The dynamic programming algorithm is a way to compare possible combinations of these sums of squared residuals to achieve a minimum global sum of squared residuals.

Thus, the first step in Bai and Perron's algorithm is to build the triangle matrix of the Sum of Squared Residuals (SSR). In the pure structural break model ($p=0$), the estimates $S_T(T_1, T_2, \ldots, T_m), \delta^*$ and $\mu_t^*$ are obtained using ordinary least squares on each segment of each possible partition. In this case, the elements of the triangle matrix SSR are calculated based on the sum of squared residuals obtained by applying least-squares to each segment $a_{ij}$, and on an updated formula to calculate recursive residuals. After that, a dynamic programming approach is used to evaluate which partition achieves a global minimization of the overall sum of squared residuals. This procedure is a sequential examination of optimal one-break partitions and contains $M$ steps. Initially, all possible one-break partitions are evaluated and a set of optimal one-break partitions (for all subsamples that allow a possible break) with their related sum of squared residuals are stored. After that, the same is done for all possible two-break partitions and so on, until a set for the $M$-break partitions is built. In each step, the breaks that produced the lowest sum of the squared residuals in the previous step are considered. Finally, a single optimal $m$-partition that produced the lowest sum of the squared residuals is determined.

In the partial structural break model ($p>0$), the dynamic programming method cannot be applied directly. The authors consider, for this case, a recursive procedure along with a method to choose the initial value for $\beta$ (to avoid local minimum). The sum of the squared residuals is written as a function of $(\beta, \theta)$, i.e., $SSR(\beta, \theta)$, where $\theta = (\delta, T_1, T_2, \ldots, T_m)$ and minimized with respect to $\theta$ keep-
ing $\beta$ fixed, next, minimized with respect to $\beta$ keeping $\theta$ fixed and then, iterated.

The test statistics and information criteria above are tools proposed in Bai and Perron (1998b) to search and to make inference on the results generated by the algorithm above.

a) Test of no break versus an unknown number of structural breaks
- UD max $F_T^*(m, q)$.
  Ho: no structural breaks
  Ha: an unknown number of structural breaks (maximum $m$ breaks).

b) Test of no break versus a fixed number, $k$, of structural breaks
- Sup $F_T(k, q)$.
  Ho: $\delta_1 = \delta_2 = \delta_3 = \ldots = \delta_{k+1}$
  Ha: $\delta_j \neq \delta_{j+1}$ for some $j$ of the given partition $k$.

c) Test of $k+1$ versus $k$ structural breaks – Sup $FT (k + 1|k)$.
  This test carries out $k + 1$ tests with Ho: no structural breaks versus Ha: one structural break, that is, $k = 1$ in the previous test. The test is applied to each segment containing observations $T_{i-1}$ to $T_i$, for $i$ from 1 to $k + 1$.

d) Information criteria
  There are two information criteria to select the dimension of a model: BIC and LWZ\textsuperscript{13}. According to Bai and Perron(1998), when there is no serial correlation in the errors, both criteria perform reasonably well. However, when serial correlation in the errors is

\textsuperscript{13}LWZ criteria are defined by

\[
LWZ = 1n \left( \mathcal{S}_T \left( T_1^*, \ldots, T_m^* \right) / (T - p^*) \right) + (p^* / T) 0.299 \left( 1n \left( T \right) \right)^{2.01}.
\]
present, they have a bias towards a higher number of structural breaks than the true value. We decided to use the BIC criteria, defined below, once it is the most efficient in the presence of structural breaks:

\[
\text{BIC}(m) = \ln(m) + p^* \ln \sigma^2(m) T^{-1},
\]

where \( p^* = (m+1)q + m + p \) and \( \sigma^2(m) = T^{-1} S_T \left(T_1^*, \ldots, T_m^*\right) \).

5. Credit and inflation in Brazil: Application of Bai and Perron's test.

Given the great number of possibilities for assessing the effects of the change in inflation \( \Delta \pi_t \) on credit \( y_t \), we considered four linear models with explanatory variables \( 1, t \) and \( \pi_t \), where \( t \) is a time trend (equations 6).

\[
\begin{align*}
y_t &= \alpha_1 + \beta_1 t + \gamma_1 \Delta \pi_t \quad & t \leq T_1 \\
y_t &= \alpha_2 + \beta_2 t + \gamma_2 \Delta \pi_t \quad & T_1 < t \leq T_2 \\
&\vdots \\
y_t &= \alpha_m + \beta_m t + \gamma_m \Delta \pi_t \quad & T_m < t
\end{align*}
\]

Table 4 shows which coefficients, \( \alpha, \beta \) or \( \gamma \), for each linear model (case), are subject to variations between structural breaks.
### Table 4
Summary of linear models

<table>
<thead>
<tr>
<th>Case</th>
<th>$\beta$</th>
<th>$\gamma$</th>
<th>$\alpha$</th>
<th>$\beta$</th>
<th>$\gamma$</th>
</tr>
</thead>
<tbody>
<tr>
<td>1</td>
<td></td>
<td>X</td>
<td>X</td>
<td>X</td>
<td>Pure model ($p=0$, i.e., $x_t={0}$ and $z_t={1, t, \Delta \pi_t}$). All coefficients can vary between structural breaks.</td>
</tr>
<tr>
<td>2</td>
<td></td>
<td>X</td>
<td>X</td>
<td>X</td>
<td>Partial model ($p=1$, i.e., $x_t={\Delta \pi_t}$ and $z_t={1, t}$). All coefficients can vary except $\gamma$.</td>
</tr>
<tr>
<td>3</td>
<td>X</td>
<td>X</td>
<td>X</td>
<td>Pure model ($p=2$, i.e., $x_t={t, \Delta \pi_t}$ and $z_t={1}$). Only coefficients $\alpha$ are subject to changes.</td>
<td></td>
</tr>
<tr>
<td>4</td>
<td>X</td>
<td>X</td>
<td></td>
<td>X</td>
<td>Pure model ($p=0$, i.e., $x_t={t}$ and $z_t={1, \Delta \pi_t}$). All coefficients can vary except $\beta$.</td>
</tr>
</tbody>
</table>

Note: The general specification is given by the set of equations (6).

In all cases:

a) serial correlation and heteroskedasticity are allowed;

b) $x_t$ is the set of explanatory variables whose coefficients are not subject to changes and $z_t$ is the set of explanatory variables whose coefficients are subject to changes;

c) We allowed $\alpha$ to vary between structural breaks. Graph 2 shows that any model that uses credit as the dependent variable should allow the mean to change over time. Thus, the set of explanatory
variables whose coefficients are subject to changes, $z_t$, always includes the constant (variable 1).

d) We used a maximum number of breaks equal to 5, since the sample size does not allow for a greater number of breaks. However, as we will see shortly, this was not a problem once the best models indicated a maximum of 4 breaks.

The results of the different cases were the following:

1) Case 1: pure model with intercept, trend and inflation\(^{14}\)

The absence of breaks is rejected by the test of no structural breaks versus an unknown number of structural breaks. The UDmax test rejected the null hypothesis at 1% (UDmax = 7.411.01).

Similarly to the previous test, the test of no structural breaks versus a fixed number, $k (k \geq 5)$, rejected the hypothesis of no structural breaks. The results of supFT\((k)\) were significant for all values of $k$, at 1% (supFT\((1)\) = 155.24; supFT\((2)\) = 94.84; supFT\((3)\) = 647.76; supFT\((4)\) = 741.10; supFT\((5)\) = 657.89).

The test of $k + 1$ versus $k$ structural breaks, given by supFT\((k + 1 | k)\), rejected, at 1%, the possibility of one structural break in favor of 2 structural breaks (supFT\((2|1)\) = 30.73), 2 in favor of 3 (supFT\((3|2)\) = 474.71) and 3 in favor of 4 (supFT\((4|3)\) = 30.73). However, it did not reject the possibility of 4 in favor of 5 (supFT\((5|4)\) = 7.00) and, thus, we concluded that the series presents 4 statistically significant structural breaks.

The sequential procedure also indicated, at 1%, that there are 4 structural breaks, reinforcing the result obtained above. The first structural break occurred in 1978:3, the second one in 1985:3, the

\(^{14}\)The results below were obtained using the IPC-Fipe to measure inflation. However, the qualitative results do not change if the IGP-DI (General Price Index - Internal Availability) or the series of official inflation rates of each period are used. We choose the IPC-Fipe because, as observed before, it has never been purged during the stabilization plans.
third one in 1989:4 and the fourth one in 1994:2. On the other hand, the BIC information criteria pointed to only 3 structural breaks (the lowest value of BIC was obtained when k equals 3). These breaks corresponded to 1985:3, 1989:4 and 1994:2. Hence, the importance of the structural change in 1978:3 is weakened, while the importance of the other three breaks (prior to the Cruzado, Collor I and Real plans) is strengthened. As the model with 3 breaks had an $R^2$ (86.8%) smaller than the model with four breaks (88%) and the highest sum of squared residuals when all the possible breaks were considered, we concluded for the model with 4 structural breaks.

The results of the estimations for the pure model with 4 structural breaks are shown below:

\[
\begin{align*}
\text{\( Y_t = 0.1092 + 0.002t - 0.0578 \Delta \pi_t \) } & \quad t \leq 1978:3 \quad (0.0050) \quad (0.0002) \quad (0.0840) \quad (1977:3 - 1979:1) \\
\text{\( Y_t = 0.0748 + 0.0001t + 0.0387 \Delta \pi_t \) } & \quad 1978:3 < t \leq 1985:3 \quad (0.0173) \quad (0.0003) \quad (0.0562) \quad (1984:1 - 1985:4) \\
\text{\( Y_t = -0.2231 + 0.0051 \Delta \pi_t \) } & \quad 1985:3 < t \leq 1989:4 \quad (0.0600) \quad (0.0008) \quad (0.0217) \quad (1989:2 - 1990:2) \\
\text{\( Y_t = -0.2966 + 0.0047 \Delta \pi_t \) } & \quad 1989:4 < t \leq 1994:2 \quad (0.0639) \quad (0.0007) \quad (0.0140) \quad (1993:4 - 1994:3) \\
\text{\( Y_t = 0.0883 + 0.009t - 0.0817 \Delta \pi_t \) } & \quad 1994:2 < t \quad (0.0866) \quad (0.0008) \quad (0.0232) \quad (0.0817) \quad (0.008) \\
\end{align*}
\]

Note: ++ and +++ mean that the coefficients are significantly different from zero at the level of significance of 5% and 1% respectively.
Graph 3 shows the credit series and the 4 statistically significant breaks, $T_1$ (1978:3), $T_2$ (1985:3), $T_3$ (1989:4) and $T_4$ (1994:2) obtained for case 1.

The four breaks coincide with important policy changes in the Brazilian economy. Indeed, these breaks occur before these great changes, and seem to indicate that somehow the credit market foresaw the most remarkable economic changes the country has gone through.

The first break (1978:3) occurred just before the second oil price shock and the international interest crisis in the late 1970s. The coefficients of the slope ($\beta_1$ and $\beta_2$) and the coefficients of $\Delta \pi_t$ ($\gamma_1$
and \( \gamma_2 \) were not significant, while coefficients of the intercept (\( \alpha_1 \) and \( \alpha_2 \)) were. As expected there was a reduction in the intercept of the trend function from 0.1092 to 0.0748.

The second structural break occurred on the eve of the Cruzado Plan (1985:3), with a nearly 40\% monthly inflation rate. The coefficients of \( \Delta \pi_t \ (\gamma_3) \) became significant only after this structural change and they were also negative, indicating a negative relationship between variation of inflation and credit. The trend function suffered a reduction in the intercept but an increase in the slope. The positive impact of the Cruzado Plan on the credit can be explained by the negative real interest rates observed after the adoption of the stabilization plan, by the real wage gains and by the price cap established by the inflation control policy.

Between the Summer Plan (the last of the heterodox plans) and Collor I Plan, the Brazilian economy moved towards a hyperinflationary process. The high level of uncertainty about the economy was enough to generate a third statistically significant structural break (1989:4). The intercept (\( \alpha_4 \)) was the lowest coefficient observed in the model and there was a reduction in the slope (\( \beta_4 \)). Both changes confirm the negative effect on the credit due to the restrictive monetary policy adopted at that time. Approximately 70\% of the financial assets were blocked and converted into fixed-term deposits under the control of the Central Bank, which obviously affected credit supply negatively. However, part of the negative effect of the plan was eliminated by the reduction in inflation once the negative and statistically significant relationship (\( \gamma_4 \)) between variation of the inflation and credit is confirmed.

The fourth and last structural break occurred at the time of
implementation of the Real Plan (1994:2). The coefficients of the intercept and slope ($\alpha_5$ and $\beta_5$) were not statistically significant. However, the coefficient of $\Delta \pi_t$ ($\gamma_5$) had the highest absolute value. Implying that inflation, in this case, caused a greater impact on credit than in the previous ones. Indeed, the remarkable increase in aggregated demand after the Real Plan led the government to adopt restrictive credit measures\textsuperscript{15}. The improvement of credit conditions (for example, better maturity dates), associated with the optimism about the economic situation, probably explains the expansion in consumption during the first years of the Real Plan.

The Bresser, Summer, and Collor II Plans did not seem to be statistically significant, which means, they did not represent fundamental changes in the inflation-credit relationship. The short duration of these plans is possibly one of the reasons for the unchanged expectations regarding credit.

2) Cases 2, 3 and 4: existence of at least one fixed coefficient.

The results obtained using the tests that fix at least one coefficient (Table 5) were qualitatively and quantitatively worse than those obtained in case 1. Nevertheless, they helped us to understand what we lose by fixing a certain variable and to perceive what did not change even when almost all coefficients were considered fixed.

Table 5

<table>
<thead>
<tr>
<th>Case</th>
<th>fixed (xt)</th>
<th>variable (zt)</th>
<th>sequential model</th>
<th>breaks</th>
</tr>
</thead>
</table>

All tests indicated the presence of at least one structural break. The tests that show more than one break had an $R^2$ greater than or equal to 81% and always included the following dates 1985:3, 1989:4 and 1994:1, occasionally considering 1978:3.

When only the intercept of the trend function was variable (case 3), the most appropriate model had only one structural break, in 1978. Despite the model having shown an $R^2$ below 70% and a confidence interval of the structural break dates with a span of up to five years, this result confirms the importance of the oil crisis for the intercept of the credit trend function. In this model, however, the coefficient for variation of the inflation, albeit negative, was not statistically significant.

When the intercept and the slope of the trend function were variable (case 2), the most suitable model had four structural breaks: 1978:3, 1985:3, 1989:4 and 1994:1. The coefficient of the change in inflation was negative and statistically significant at 1%.

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$^{16}$When we considered the models with no structural breaks, the value of $R^2$ drops to 53%, confirming that the models lose information when structural changes are not considered.

$^{17}$The results of the estimations are available from the authors upon request.
In case 4, when only the slope of the trend function was fixed, the most adequate model had four structural breaks: 1978:3, 1985:4, 1990:1 and 1994:2. The coefficient of the change in inflation was negative and became statistically significant at 1%, only after the third structural break. However, the model had a confidence interval that comprised up to seven years, which hinders the analysis of the results. In summary, as we also observed in case 3, imposing a fixed slope results in less reliable confidence intervals.

6. Credit, inflation and interest rates in Brazil: Application of the Bai and Perron test.

Obviously inflation is not the only variable that affects credit. Therefore, in order to test the robustness of our results, we included the real interest rate as another explanatory variable. The interest rates are Selic rates and cover the first quarter of 1974 (1974:1) to the last quarter of 1998 (1998:4). The traditional ADF test rejected the unit root hypothesis, indicating that the series is stationary\(^{18}\).

We initially considered the pure linear model (all coefficients are subject to changes between breaks) with the following explanatory variables 1, \(t\), \(\Delta \pi_t\), \(r_t\), where \(r_t\) is the real interest rate. As in section 5, we allowed for serial correlation and heteroskedasticity and we established a maximum of 5 breaks.

The test of no structural breaks versus an unknown number of structural breaks, given by the Udmax, rejected the null hypothesis at 1% (UDmax = 846.14). Likewise, the test of no structural breaks versus a fixed number, \(k (k \leq 5)\), given by sup\(FT\), rejected the absence of structural breaks. The results of sup\(FT(k)\) were significant for all the values of \(k\), at 1% (sup\(FT(1) = 246.41; sup\(FT(2)\)

\(^{18}\)Values available from the authors.
Maria Beatriz Zerbini and Fabiana Rocha

\[ = 249.30; \text{sup}FT(3) = 797.99; \text{sup}FT(4) = 716.67; \text{sup}FT(5) = 846.14\).

After the two tests above have rejected the hypothesis of no break, the test of \( k + 1 \) structural break versus \( k \) structural break concluded that the series presents the maximum number of breaks, which is 5. The \( \text{sup}FT(k + 1|k) \) test rejected, at 1%, the possibility of one structural break in favor of 2 (\( \text{sup}FT(2|1) = 67.51 \)), 2 in favor of 3 (\( \text{sup}FT(3|2) = 166.50 \)), 3 in favor of 4 (\( \text{sup}FT(4|3) = 22.57 \)) and 4 in favor of 5 (\( \text{sup}FT(5|4) = 34.82 \)).

However, the information criteria (BIC) and the sequential procedure indicated the presence of 3 structural breaks at 1%. Therefore, we chose the model with 3 structural breaks.

The results for the sequential procedure model with \( R^2 \) equal to 89.2% are presented below:

\[
\begin{align*}
 y_t &= 0.1110^{+++} -0.0008^{+++}t +0.089\Delta \pi_t +0.1549\pi_t & t \leq 1985:3^{++} \\
 y_t &= -0.0304^{(0.0392)} +0.003^{+++}t -0.0016\Delta \pi_t +0.1215\pi_t & 1985:3 < t \leq 1990:3^{(0.0292)} \\
 y_t &= -0.189^{+++}(0.0671) +0.0043^{+++}t -0.0271\Delta \pi_t -0.079^{+++}\pi_t & 1990:1 < t \leq 1994:3^{(0.054)} \\
 y_t &= 0.0582^{(0.0791)} +0.0013t -0.0853^{+++}\Delta \pi_t +0.173^{+++}\pi_t & 1994:1 < t
\end{align*}
\]

Note: +, ++, and +++ mean that the coefficients are statistically different from zero at the 10%, 5% and 1% levels, respectively.

In general, this model brought little additional information about the relationship between inflation and credit. Basically, the trend function explains the credit behavior up to the Collor Plan when the coefficients of the change in inflation and the real interest rates become statistically significant.

As in case 1 from subsection 5, the inflation coefficient was negative. However, real interest rates showed a positive coefficient after
the Real Plan\textsuperscript{19}. The results were contrary to the hypothesis that real interest rates and credit go in different directions. However, the literature\textsuperscript{20} shows that, at the initial stages of the stabilization plans, mainly of heterodox ones, the reduction in inflation is higher than the reduction in nominal interest rates, which results in higher real interest rates. In the Real Plan, the decrease in inflation and the perspective of a favorable future economic environment contributed to the improvement of credit despite the increase in the real interest rates in the short term.

7. Conclusions.

The purpose of this paper is to investigate the relationship between credit, inflation, and stabilization plans in Brazil during 1970-1998.

After determining the order of integration of credit and inflation, we used Bai and Perron’s methodology (1998b) to endogenously determine the structural changes to the credit series. Several models were considered, from a pure structural change model where all the coefficients are subject to changes up to a model where only the intercept of the trend function varies. The pure model brought a richer analysis as it allowed us to evaluate the behavior of inflation in each of the break periods of the deterministic trend.

The hypothesis of no structural breaks was rejected for all models. Besides, the models with two or more breaks presented a higher $R^2$ than the models with no structural breaks. Therefore, information is lost when the possibility of structural changes is not considered.

\textsuperscript{19}The results for cases 2, 3 and 4, where at least one coefficient is fixed, when the interest rate is included, can be directly obtained from the authors. As in cases 2, 3 and 4 from section 5, they are qualitatively and quantitatively worse than those obtained for case 1.

\textsuperscript{20}Khamis (1996).
Statistically significant breaks always included the following periods: 1985:3, 1989:4 and 1994:1 and occasionally 1978:3, which means that the breaks took place during the plans or in the period immediately before them. It seems that somehow the credit market anticipated the economic changes.

The models captured the effects of the second oil price shock and the international interest crisis in the late 1970s and the positive effects of the Cruzado Plan due to the reduction in inflation and to specific characteristics of the plan –negative real interest rates, real wage gains and price caps generated by the anti-inflationary policy. Indeed, the coefficient for the variation of inflation becomes statistically significant only after the adoption of this plan. The models also captured the negative effects of the restrictive monetary policy adopted during Collor I Plan. After this plan the intercept of the trend function reached the lowest value and there was a reduction in its slope. Finally, the models showed that the change in inflation generated a higher impact on credit during the Real Plan than in the previous plans. The belief on the success of the Plan apparently helped to change credit behavior, despite the high interest rates observed after its implementation.

All results indicated a negative and statistically significant relationship between credit and the change in inflation. We could conclude, then, that the stabilization plans contributed to credit expansion. The question that remains is how to promote only good quality credit. The adoption of prudential policies with the aim of improving the risk evaluation is, therefore, deemed necessary in spite of the challenge in avoiding extremely restrictive regulations.

Private sector credit and inflation during Brazilian stabilization plans

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