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A cointegration analysis of crime, economic activity, and police performance in São Paulo city

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The main objective of this paper is to investigate possible causes for the significant reduction observed in crime rates in São Paulo city. By applying a cointegration analysis, we observed long-run relationships between crime, economic activity, and police performance. The results indicate that the lethal crime rate is positively related to unemployment and negatively related to real wages and to the results of law-enforcement activities, specifically arrests and seizure of firearms. Moreover, the hypothesis that the Disarmament Statute led to a reduction in the lethal crime rate is not rejected.

Keywords: crime; time series; cointegration analysis; criminality; lethal crime

1. Introduction

The city of São Paulo (hereinafter just São Paulo) managed to interrupt and reverse a marked trend toward rising crime rates, particularly lethal crime rates. This was a major achievement, which led São Paulo to being ranked among a select group of cities with similar successful experiences in this area, particularly New York and Bogota.

In this context, a challenging question came up: what is the cause or causes of the reduction in crime rates in São Paulo? No answer supported by empirical evidence is available so far. Some experts believe that this reduction resulted from the disarmament policy implemented in the state since 2001 and was enhanced by the passage of the Disarmament Statute (DS)[1], while others claim that law enforcement has become more active and efficient, and some others believe that it was brought about by better economic conditions, such as a significant drop in unemployment. However, the only evidence available is that reported by Cerqueira [12] and by Santos and Kassouf [77]. Although they were not intended to answer the question under discussion, those studies shed the first light on the subject. Souza *et al.* [83], Goertzel and Kahn [30], and Peres *et al.* [67] also

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discussed the reduction in the crime rate in São Paulo from a qualitative point of view, but very appropriately.

In Santos and Kassouf [77], the aim was to evaluate the effect of the DS on lethal crime rates in São Paulo. The hypothesis that the disarmament policy led to a reduction in this crime indicator was not rejected. Although plausible, this was a conclusion based on results obtained by intervention analysis, where the information used in the empirical modeling was drawn from the crime rate itself. Nevertheless, it corroborates the conclusion of Cerqueira [12] on the causal relationship between firearms and crime in municipalities of São Paulo state. In this study, a proxy variable was inevitably used for the amount of firearms in circulation to circumvent the endogeneity of this variable, and it was assumed that the under-reporting of crimes of all types that were analyzed was invariant over the analyzed period. In the São Paulo case, Santos and Kassouf [77] argue against this hypothesis for some crimes, especially crimes against property. We will resume this discussion in Section 3.

In this context, the main purpose of this study is to identify, in greater detail, the causes of the significant reduction observed in crime rates in São Paulo. Based on previous evidence presented by Santos and Kassouf [77], another statistical methodology was applied in order to allow for other alleged determinants of crime reduction to be included in the empirical model specification. Specifically, we were looking for evidence for the assumptions made in the second paragraph of this section.

It should be noted that most economic studies on the causes of crime are based on panel aggregate data, especially in Brazil. In this paper, the problem is addressed by analyzing data from a single spatial unit by means of a time-series econometric methodology. This approach avoids the bias due to spatial dependence between municipalities or even between states. It also reduces the possibility of bias due to measurement errors resulting from differences in data collection procedures and quality between different locations. It should also be noted that the methodology does not require the assumption of exogeneity for any of the variables, so it was not necessary to use instrumental variables for potentially endogenous variables. A last distinction to be made is that it was possible to use indicators for the results of police activities instead of public security spending, which is a variable commonly used to reflect these results.

It should also be noted that few studies, even in the international literature, analyze the causes of crime using recent advances in time-series econometrics. In this regard, special mention should be made to the studies by Corman and Mocan [16], Saridakis [78] and, recently, Saridakis [79]. The first study was conducted with monthly data for New York, the second one used annual data for the USA, and the last one was based on data for England and Wales. Therefore, we are enriching this literature by presenting specific evidence for São Paulo, where the problem can be analyzed from a different perspective: that of crime reduction.

After this introduction, the structure of the statistical model will be presented in Section 2; details about the sample, variables and data are provided in Section 3; the specification of the empirical model is provided in Section 4; and results are discussed and conclusions presented in Sections 5 and 6, respectively.

2. Structure of the model

The starting point of the empiric modeling is the basic form of an autoregressive vector model (VAR) [82], consisting in a set of K endogenous variables $\mathbf{y}_t = (y_{1t}, \dots, y_{kt}, \dots, y_{Kt})$. The VAR($p$) process is defined by:

$$\mathbf{y}_t = \mathbf{A}_1\mathbf{y}_{t-1} + \cdots + \mathbf{A}_p\mathbf{y}_{t-p} + \mathbf{u}_t, \tag{1}$$

where $\mathbf{A}_i$ are the $(K \times K)$ coefficient matrices for $i = 1, \ldots, p$ and $\mathbf{u}_t$ is a K-dimensional process with $E(\mathbf{u}_t) = 0$ and time-invariant positive definite covariance matrix $E(\mathbf{u}_t, \mathbf{u}_t') = \mathbf{\Sigma}_u$ (white noise).

In addition to the variables within the autoregressive vector, relying on the diagnosis made by Santos and Kassouf [77], Equation (1) is expanded to include an intervention dummy variable for the DS. Centered seasonal dummy variables (S) were also added for seasonality. So the initial specification of the model is as follows:

$$\mathbf{y}_t = \mathbf{A}_1\mathbf{y}_{t-1} + \cdots + \mathbf{A}_p\mathbf{y}_{t-p} + \Phi\mathbf{D}_t + \mathbf{u}_t, \tag{2}$$

where $\mathbf{D}_t$ is the matrix composed of the centered seasonal dummies and of the intervention dummy, where DS $= 0$ for $t < 2004Q1$ and DS $= 1$ for $t \geq 2004Q1$.[2]

Other specific details about the methodology and hypothesis testing will be provided below, along with the results.[3]

3. Sample, variables and data

The sample used in the estimations is composed of 56 observations for São Paulo between the first quarter of 1997 and the fourth quarter of 2010.

The model specification given by Equation (2) is composed of three types of indicators: crime, economic activity, and police performance. Specifically, the effects of the two latter indicators on crime are analyzed. It is a variant of the crime supply curve proposed by Becker [6] and Ehrlich [21], among others.

The level of economic activity is a proxy for both opportunity costs of crime and for the expected return on it. The net effect of its effect can be either positive or negative, meaning that it is a question that can only be answered empirically. It is expected, however, that crime is negatively related to wages and positively related to unemployment. Specifically, it is expected that better conditions in the labor market cause crime levels to drop.

The last indicator is a proxy for variables that, in theory, have a deterrent effect on criminal behavior. Their effects on crime are expected to be negative.

At this point, an observation should be made about the quality of the crime indicator. Since what is being investigated is whether there is any relationship between economic activity and crime, among other possible relationships, the first impression is that only crimes against property, i.e. crimes with strictly economic motivations, should be considered. Although analyses such as this one are common in empirical studies, particularly in those conducted with US data, the robustness of the estimates can be challenged. According to MacDonald [58], discussions on the relationship between economic cycles and crime have been restricted to the econometric methodologies applied, without any mention of the quality of the analyzed crime indicators. The author shows empirically that the rate of under-reporting of crimes against property is sensitive to economic conditions, but in the opposite direction to that of the sensitivity of crime itself. In other words, crime under-reporting rates rises during economic downturns and decreases during periods of economic growth. If this is true, the rate of under-reporting of crimes against property decreased in São Paulo during the period studied here, as growth in employment clearly shows that high economic growth was experienced during that period. For this reason, and for all the other reasons provided by Santos and Kassouf [77], we decided to use the rate of lethal crimes – homicide and robbery aggravated by death – per one hundred thousand population (crime) as the crime level indicator. The quarterly lethal crime rate per one hundred thousand population (hereinafter just crime) was calculated by interpolating data from annual estimates of residing population published by the Brazilian Institute for Geography and Statistics.

The performance of law enforcement in preventing and fighting crime is measured by means of two measures built from indicators of direct results (firearms seized, reported drug trafficking crimes and total arrests), which vary according to the crime rate, and indicators of indirect activity (searches or identification of persons, such as, for instance, blitzes), which reflect what law enforcement is doing [11]. These are police performance indicators that, as mentioned above, are used for building proxies for variables causing deterrent effects on criminal behavior by increasing the probability of crime failure.

Selecting 'good' performance indicators for law enforcement is a hard task. And, as in the case of the crime indicator, results depend on the quality of these indicators. Arguments that justify our choice will be briefly presented below.

The number of firearms seized by police is an indicator that can be interpreted both as a proxy for the amount of weapons in circulation and as an indicator of the results of police activity [49]. For the first interpretation, it is assumed that fewer firearms are seized because there are fewer guns in circulation on the streets, and for the second one it is assumed that more guns are seized because law enforcement is carrying out more operations to remove them from circulation. However, it must be recognized that most weapons are seized during operations designed to fight other crimes, such as drug trafficking. Therefore, they vary according to the amount of crime.[4] In time-series analyses, it should be stressed that it only makes sense to say that the absolute measure of seized weapons reflects the amount of weapons in circulation if and only if police productivity is constant over time. This hypothesis is too strong, as there were changes in the police 'technology' used in São Paulo state and in the city of São Paulo in the period considered in this study. These advances in the technology used by law enforcement may have affected its performance.

It is also obvious that inspection, law enforcement, and disarmament campaigns held after the DS was regulated secured an unquestionably positive result: they reduced the number of firearms in circulation. According to a victimization survey carried out in São Paulo in 2003 and 2008 [17], the number of people with a firearm at home decreased by 11.5% over that period; it also showed that the use of firearms in residence burglaries and thefts of persons decreased by 69% and 14.2%, respectively. With fewer guns in circulation, weapons are, ceteris paribus, less liable to seizure. In fact, data produced by the Public Security Secretariat of São Paulo show that the number of firearms seized in the city over the same period declined by 51.2%.

In this context, for this variable to be seen as an indicator of police performance, it needs to be measured in relation to what law enforcement actually does. For this purpose, the variable defined by the ratio between the number of seized firearms and searches and identification of persons by police officers (guns) is used. Thus, the variable built from the ratio between these two measures reflects a relative number of seized firearms.

It should be observed that the relative amount of searches and identification of persons increased significantly. The rate of searches of persons per one hundred population rose from 1.98 to 5.83 over the period. Although more operations of this kind might have been conducted with the same number of police officers and equipment, it is more likely that this increase was a direct result of the

Table 1. Series definitions and summary statistics, São Paulo, 1997Q1–2010Q4.

Series	Definition	Mean	Std. Dev.	Min.	Max.
crime	Crime rate per one hundred thousand population	8.48	4.41	2.44	14.62
arrest	Percentage of arrests for drug trafficking	14.43	5.99	6.19	25.32
guns	Amount of firearms seized per one hundred thousand searches or identification of persons	756.38	465.37	185.41	1677.46
unemp	Total unemployment rate (%)	15.57	2.31	10.50	19.57
wage	Real minimum wage (in Brazilian Reals)	389.76	87.02	279.28	556.15

hiring of more police personnel and purchase of more equipment, such as vehicles and weapons. It is therefore assumed that increases in this variable directly reflect the increased presence of law enforcement on the streets and subsequent higher probability of crime failure and, consequently, of crime being deterred.

The second police performance indicator was proposed to reflect the fight against drug trafficking specifically. For this purpose, the ratio between the number of drug trafficking crimes[5] and total arrests is used as a measure (arrest). Since the first variable measures the number of occurrences and not seized quantities and the second one measures the number of arrests and not of people arrested, it is assumed that an arrest occurs for every drug trafficking crime that is reported. The variable built from the ratio between these two measures is therefore an approximation of the percentage of total arrests of drug dealers.

The data used for building the three variables related to public security are the ones published in the Official Gazette of the State of São Paulo by the Public Security Secretariat of São Paulo (SSP-SP).[6]

We used two measures to reflect the local economic activity level: total unemployment rate (unemp)[7] and real minimum wage (wage). It should be noted that both variables are measures

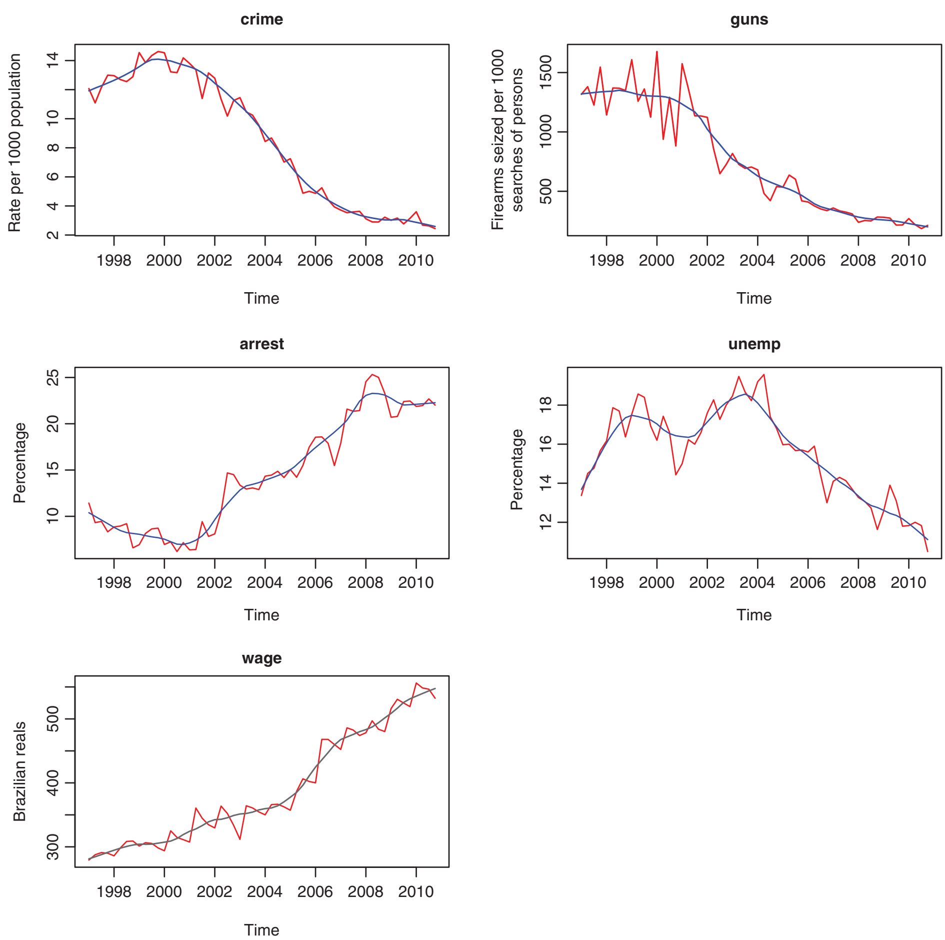

Figure 1. Series crime, guns, arrest, unemp and wage – São Paulo, 1997Q1–2010Q4.
Source: Prepared by the authors. The series definitions are provided in Table 1.

specifically applied to São Paulo that were respectively provided by the State Data Analysis System Foundation (SEADE)[8] and by the Inter-Union Department of Statistics and Socioeconomic Studies (DIEESE).[9]

Table 1 shows summary statistics of the data set that was used. Figure 1 shows the time paths of the series during the sampling period. To facilitate viewing of possible time trends in the series, a Lowess smoothing line [14,15] was included in each of their graphs.

4. Model identification

4.1 *Data transformation*

The model estimation expressed by Equation (2) requires that all series are stationary. This property will be checked in the next subsection. It is known, however, that the presence of a stochastic trend can be associated to the behavior of the data variance. If this is the case, prior to applying unit root tests, it is necessary to stabilize the series variance by applying a suitable transformation to the data.

The need for transformation was assessed by means of Box–Cox analyses [9]. This procedure consisted in estimating, by maximum likelihood, transformation parameter λ in the family of transformations $y_t^* = (y^\lambda - 1)/\lambda$ if $\lambda \neq 0$ and $y_t^* = \log(y_t)$ if $\lambda = 0, t = 1, \ldots, T$. Assuming that the data are not i.i.d., it was concluded that two of the five series did not require transformation. For one of them, however, logarithmic transformation is not rejected at 95% confidence level. In the case of the three other series, which require transformation, logarithmic transformation is also appropriate. In this context, the conventional approach of applying logarithmic transformation to all series is applied. Therefore, the logarithm was taken for all variables before unit root tests were applied and, as appropriate, before taking the first differences of the series. The time paths of the logarithms of the series are shown in Figure 2.

4.2 *Unit root tests*

In economic studies, it is a very usual procedure to clearly indicate that macroeconomic series, such as unemployment and wages, constitute non-stationary processes. In public security series, however, very little is known about the properties of the stochastic process generating the data. Thus, apart from essential to the empirical modeling that will be presented in the following sections, this section provides new evidence and constitutes a marginal contribution to this study, particularly to future empirical analyses of the causes of crime using time series data.

Prior evidence suggests that crime indicators are non-stationary variables [34,72,79]. In addition, there may be more than one unit root [72]. Given this possibility, the first step was applying the Dickey–Pantula test [18] to test for the hypothesis of two unit roots. The results are shown in Table 2.

In the first step of the test, null hypothesis H_0: $d = 2$ was tested against alternative hypothesis H_A: $d = 1$, judging the statistical significance of the estimated coefficient, $\hat{\beta}_1$, in model $\Delta^2 y_t = \alpha + \beta_1 \Delta y_{t-1} + \varepsilon_t$. The hypothesis of two unit roots is rejected in all series. In the second step, null hypothesis H_0: $d = 1$ was tested against alternative hypothesis H_A: $d = 0$, estimating model $\Delta^2 y_t = \alpha + \beta_1 \Delta y_{t-1} + \beta_2 y_{t-1} + \varepsilon_t$, and assessing the statistical significance of both coefficients, $\hat{\beta}_1$ and $\hat{\beta}_2$. For all series, the results of this step, in terms of the significance of $\hat{\beta}_1$, are not different from those of the first stage. And for all series the hypothesis of a unit root is not rejected.[10]

Once the presence of two unit roots is ruled out, the existence of a unit root in the stochastic process generating the data is evaluated by means of the ADF–GLS test [22] and KPSS test [50]. In both tests, the model specification contains a constant and a trend as deterministic regressors. In the first one, null hypothesis H_0: $d = 1$ was tested against alternative hypothesis H_A: $d = 0$,

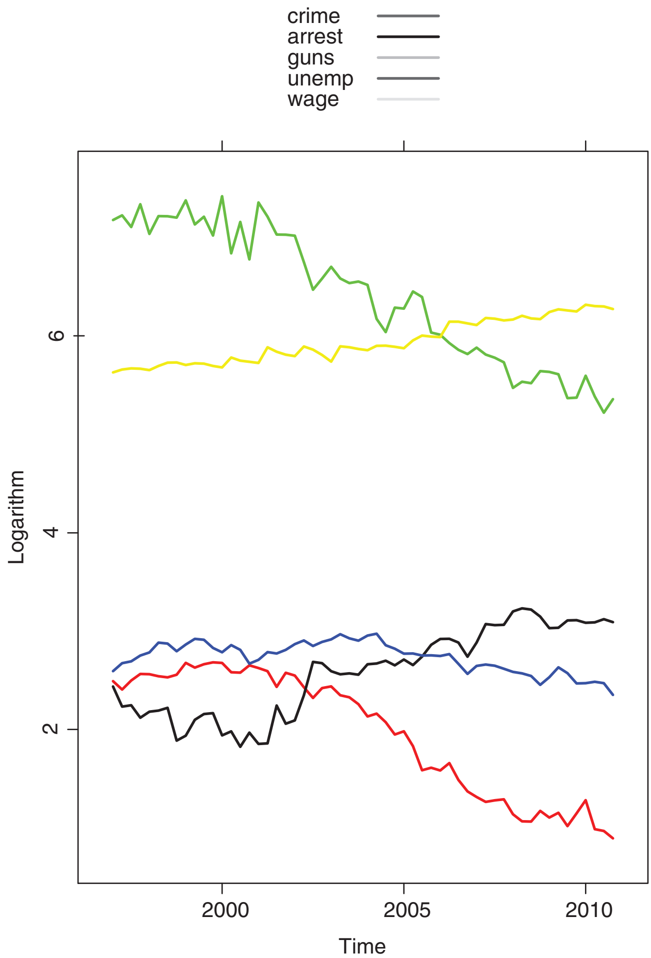

Figure 2. Logarithm of the series crime, arrest, guns, unemp and wage – São Paulo, 1997Q1–2010Q4. Source: Prepared by the authors. The series definitions are provided in Table 1.

Table 2. Dickey–Pantula test.[a]

Series	p	$\hat{\beta}_1$ e $\hat{\beta}_2$	p-Value	Conclusion
crime	1	−1.317	0.000	
	1	−1.136 and 0.016	0.000 and 0.992	I(1)
guns	1	−1.595	0.000	
	0	−1.504 and −0.00562	0.000 and 0.938	I(1)
arrest	1	−1.213	0.000	
	0	−1.029 and −0.0301	0.000 and 0.820	I(1)
unemp	1	−1.297	0.000	
	1	−1.322 and 0.0196	0.000 and 0.978	I(1)
wage	3	−1.247	0.041	
	3	−1.2303 and −0.002875	0.074 and 0.942	I(1)

Notes: [a]The first and second lines report the results of the first and second step of the test, respectively; the Mackinnon p-value is reported. The series definitions are provided in Table 1.

Table 3. ADF–GLS and KPSS tests.[a]

Series	p	Test value	Conclusion
crime	5	−1.366	
	2	0.358	I(1)
guns	4	−1.764	
	2	0.207	I(1)
arrest	1	−1.920	
	2	0.173	I(1)
unemp	2	−0.765	
	2	0.384	I(1)
wage	4	−2.110	
	2	0.305	I(1)

Notes: [a]The first and second lines report the results of the ADF–GLS and KPSS tests, respectively; series definitions are provided in Table 1.

with critical values at 1% and 5% equal to −3.58 and −3.03, respectively. Conversely, the second one tested hypothesis H_0: $d = 0$ against hypothesis H_A: $d = 1$, with critical values for the same significance levels equal to 0.146 and 0.216, respectively.[11]

The results are shown in Table 3. Comparing the critical values to the test statistics, it can be concluded that, for all series, the results of the second test corroborate those of the first one, namely, that all series have a unit root.

Note, however, that Santos and Kassouf [77] did not reject the hypothesis that the DS led to a reduction in the lethal crime rate, i.e. in the crime series. In the presence of structural breaks, conventional tests are biased toward not rejecting the unit root hypothesis. For this reason, with the aim of checking the conclusions drawn from the three previous tests, the exogenous structural break test proposed by Perron [68] was applied. This test was applied by estimating the following model, with a constant and a trend:

$$\text{crime}_t = a_0 + \mu_1 D_P + \mu_2 D_L + a_2 t + a_1 \text{crime}_{t-1} + \sum_{i=1}^{k} \beta_i \Delta\text{crime}_{t-i} + \varepsilon_t, \tag{3}$$

where D_P is a pulse dummy variable defined by $D_P = 1$ if $t = 2004Q1$ and 0 otherwise, D_L level dummy variable defined by $D_L = 1$ se $t \geq 2004Q1$ and 0 otherwise, ε_t is white noise. The lag lengths (i.e. the value k) were determined by applying the same criterion adopted by Perron [68]. We selected $k = 2$, since the t-statistic on β_2 was greater than 1.6 in absolute value and the t-statistic on β_3 was smaller than this value. a_0, μ_1, μ_2 and a_2, the null hypothesis is that the coefficient is equal to zero; a_1, the null hypothesis is $a_1 = 1$.

Coefficient $\hat{\mu_1} = -0.144$ is not statistically significant ($t = -1.58$) at conventional levels (i.e. until 10%), but it is significant at 13% (p-value$= 0,121$). However, $\hat{\mu_2} = -0.137$ is statistically significant ($t = -2.49$; p-value$= 0.017$). This result corroborates the kind of structural break that was diagnosed by Santos and Kassouf [77]. The null hypothesis that $a_1 = 1$ is assessed using the critical value at 5% simulated by Perron [68] for $\lambda = 29/56 \approx 0,5$, where λ is the proportion of observations before the structural break. Since the test value is $t = (0.7503 - 1)/0.06838 \approx -3.65$ and the critical value is equal to -3.76, the hypothesis of a one-time change in the level of a unit root process is not rejected.

It is plausible that the guns and arrest series also were influenced by the DS. Therefore, we decided to apply the Perron test to these series as well. Replacing only the series, we used the previous model specification (i.e. Equation 3). Using the same criterion for selecting the lag lengths, $k = 1$ and $k = 0$ were selected for these series, respectively. For both series, coefficients

$\hat{\mu_1}$ and $\hat{\mu_2}$ are not statistically significant at conventional levels. For the a_1 coefficient, the test values for these series are $t = (0.4331 - 1)/0.1616 \approx -3.51$ and $t = (0.7187 - 1)/0.0874 \approx -3.22$, respectively. Since the critical value is equal to -3.76, the hypothesis of a unit root process is not rejected once again.

Finally, it can be concluded that all time series of empirical model have a unit root, i.e. they are integrated of order one – I(1). Thus, the model expressed in Equation (2) should be estimated from the first difference in the series, Δy_t. But when a difference is applied to turn them into difference-stationary, the possibility of analyzing long-term relationships between them is lost. However, if there is at least one stationary relationship between them (i.e. if they are cointegrated), these relationships can be recovered [25,41,42,70,71].

4.3 *Cointegration analyses*

This study is not the only one attempting to identify a cointegrating relationship in systems composed of crime indicators, among other variables. Using Brazil data, however, it constitutes a novel application of Johansen's cointegration analysis [41–43,46] to identify the long-run effects of crime determinants.[12]

The presence of a single cointegration relationship, r, is sufficient to ensure the existence of a linear link between the stochastic trends in the series under analysis. Thus, if $r \geq 1$, the VAR, expressed in Equation (2), assumes the general representation of a vector error-correction model (VECM). The specification is given as follows ('transitory' form):

$$\Delta \mathbf{y}_t = \boldsymbol{\alpha}\boldsymbol{\beta}'\mathbf{y}_{t-1} + \Gamma_1 \Delta \mathbf{y}_{t-1} + \cdots + \boldsymbol{\Gamma}_{p-1}\Delta \mathbf{y}_{t-p-1} + \Phi \mathbf{D}_t + \mathbf{u}_t \tag{4}$$

with $\boldsymbol{\Gamma}_i = -(\mathbf{A}_{i+1} + \cdots + \mathbf{A}_p)$ for $i = 1, \ldots, p-1$, and $\boldsymbol{\Pi} = \boldsymbol{\alpha}\boldsymbol{\beta}' = -(\mathbf{I} - \mathbf{A}_1 - \cdots - \mathbf{A}_p)$. The dimensions of $\boldsymbol{\alpha}$ and $\boldsymbol{\beta}$ is $K \times r$, where r is the cointegration rank, i.e. how many long-run relationships between the variables $\mathbf{y}_t$ do exist. The matrix $\boldsymbol{\alpha}$ is the loading matrix and the coefficients of the long-run relationships are contained in $\boldsymbol{\beta}$.

The optimal lag length (p) for the unrestricted VAR model including a constant and a trend as deterministic regressors for a maximal lag length of four, was determined by a joint analysis of Akaike information criterion – AIC [3,4], Hannan–Quinn information criterion – HQ [36], final prediction error – FPE [1,2] and Schwarz Bayesian criterion – SBC [80]. The model specification includes a constant and a trend. The results are shown in Table 4.

All four criteria indicated $p = 1$. However, apart from being restrictive in relation to cointegration analyses, this lag length did not generate white noise residuals. For this reason, we decided to estimate and apply diagnostic tests for residuals to the set of models with $p \in \{2, 3, 4\}$. VAR(2) and VAR(3) were maintained as tentative candidates for the following cointegration analysis.

Table 4. Determination an optimal lag length for VAR for a maximal $p = 4$.

	p			
Criteria	1	2	3	4
AIC	−2.62[a]	−2.60	−2.59	−2.55
HQ	−2.54[a]	−2.48	−2.44	−2.37
SBC	−2.41[a]	−2.30	−2.20	−2.07
FPE	4.37e-12[a]	5.70e-12	7.14e-12	1.31e-11

Note: [a]Lag length selected.

Table 5. Johansen cointegration tests (trace).

Hypothesis [a]		Test statistics[b]		Critical values	
Null	Alternative	$p = 2$	$p = 3$	95%	99%
		Specification I			
$r = 0$	$r > 0$	88.84	90.25	68.56	76.07
$r \leq 1$	$r > 1$	45.10	43.73	47.21	54.46
		Specification II			
$r = 0$	$r > 0$	96.80	97.71	76.07	84.45
$r \leq 1$	$r > 1$	52.34	50.56	53.12	60.12
		Specification III			
$r = 0$	$r > 0$	102.59	115.34	87.31	96.58
$r \leq 1$	$r > 1$	58.82	67.82	62.99	70.05

Notes: [a] r is the cointegration rank.
[b] p is the lag-order of the VAR.

Regarding the inclusion of deterministic components in the cointegration model that will be tested, three specifications will be evaluated: constant in the short-run model (specification I); constant is restricted to the long-run model, i.e. the cointegration space (specification II); a restricted linear in the cointegration space and a constant in the short-run model (specification III).[13] Later, however, specific tests are applied to assess the statistical significance of the deterministic regressors in the cointegration relationship.

The Johansen cointegration tests (trace statistic, λ_{trace}) are shown in Table 5.[14] Regardless of the model specification, both in relation to the order and with inclusion of deterministic regressors in the cointegrating space, the existence of at least one cointegrating vector (i.e. $r = 1$) is not rejected at the 1% significance level. Therefore, there is strong evidence of a cointegration relationship in the time series. It is known, however, that the inclusion of dummy variables makes the critical values simulated by Osterwald-Lenum [65] inappropriate for the Johansen cointegration tests [48]. However, the distance observed between the test statistic values and their known critical values, particularly at the 5% significance level, ensures that the test conclusion would not be different if the adjusted critical values had been simulated. Nevertheless, in order to check the results of the tests that were applied previously (i.e. with critical values not corrected for structural break), which indicated a cointegration relationship, we calculated the critical values corrected for the trace tests following the procedures proposed by Johansen *et al.* [48].

The above-mentioned authors developed two variants of the usual trace test for cointegration among non-stationary time series: the $H_l(r)$ and $H_c(r)$ tests for when there are $(q - 1)$ breaks (i.e. q sub-samples) in a linear trend or in constant-level data, respectively.[15] The asymptotic distributions of the test statistics depend on the number of variables, on the value of the cointegration rank and the locations of the break-points in the sample denoted as $v_j = T_j/T$, where T is the full sample size and T_j is the last observation of the jth sub-sample ($j = 1, 2, \ldots, q$). Exact analytic expressions for the asymptotic distributions are not known and the quantiles must be calculated by simulation. Fortunately, Giles and Godwin [29] provide a code for the R statistical package [73] that enabled us to easily calculate the asymptotic critical values for these statistical tests.

In this study, there is a single break ($q = 2$). Hence, there are two relative sample lengths: $v_1 - 0$ and $1 - v_1$. Considering the date on which the DS was sanctioned in Brazil (22 December 2003), it was calculated that $v_1 = 0.5179$.

The asymptotic critical values for the 95th percentile for null hypothesis $r = 0$ against $r > 0$ are 88.76 and 113.52 for $H_c(r)$ and $H_l(r)$, respectively. And for null hypothesis $r = 1$ against $r > 1$, the values are 64.14 and 84.39, respectively. Comparing the critical values to the trace

Table 6. Diagnostic tests of VECM(2) – specification II.[a]

Univariate tests			
Series	Ljung–Box ($K = 14$, df $= 12$)	ARCH-LM ($K = 4$, df = 4)	Jarque–Bera (df = 2)
crime	$Q = 12.79$ [0.38]	$\chi^2 = 7.71$ [0.10]	$\chi^2 = 1.86$ [0.39]
arrest	$Q = 19.38$ [0.08]	$\chi^2 = 4.46$ [0.35]	$\chi^2 = 2.83$ [0.24]
guns	$Q = 13.01$ [0.37]	$\chi^2 = 1.30$ [0.86]	$\chi^2 = 0.80$ [0.67]
unemp	$Q = 12.29$ [0.43]	$\chi^2 = 1.53$ [0.82]	$\chi^2 = 1.06$ [0.59]
wage	$Q = 11.26$ [0.51]	$\chi^2 = 1.53$ [0.82]	$\chi^2 = 1.13$ [0.56]
Multivariate tests			
Portmanteau			
$K = 14$ (df $= 280$)	$Q = 303, 22$ [0.16]		
Edgerton–Shukur			
$K = 1$ (df $= 25; 90$)	$F = 0.80$ [0.73]		
$K = 2$ (df $= 50; 90$)	$F = 1.09$ [0.35]		
$K = 3$ (df $= 75; 71$)	$F = 0.86$ [0.74]		
$K = 4$ (df $= 100; 48$)	$F = 1.14$ [0.32]		
$K = 5$ (df $= 125; 24$)	$F = 0.90$ [0.66]		
ARCH-LM			
$K = 1$ (df $= 225$)	$\chi^2 = 225.39$ [0.48]		
$K = 2$ (df $= 450$)	$\chi^2 = 435.53$ [0.68]		
$K = 3$ (df $= 675$)	$\chi^2 = 677.80$ [0.46]		
$K = 4$ (df $= 900$)	$\chi^2 = 735.00$ [1.00]		
Jarque-Bera			
Skewness (df $= 5$)	2.69 [0.75]		
Kurtosis (df $= 5$)	2.85 [0.72]		
Both (df $= 10$)	5.54 [0.85]		

Notes: [a]*p*-Values between brackets; Portmanteau test adjusted to small samples; LM Breusch–Godfrey modified for small samples [20]; series definitions are provided in Table 1.

test statistics (Table 5), it can be concluded that, for both specifications (II and III), this analysis indicates that $r = 1$, which is consistent with our prior results.

It is worth mentioning that this study was not the only one that identified a cointegration in a model composed of crime indicators, among other variables. A cointegration relantionship was also identified by Corman and Mocan [16] and Saridakis [79], among other authors. But it was rejected in the cointegration analyses carried out by Scorcu and Celline [81], Hale [34] and Saridakis [78].

In order to proceed with analyzing long-run relationships between crime and other variables, one of six estimated models must be chosen. Specifically, the behavior of residuals and the statistical significance of the deterministic regressors in the cointegration vector still need to be checked.

The VECM(1) was discarded for not generating white noise residuals. Fortunately, the diagnostic tests applied to the VECM(2) residuals suggest that they are robust. The results of the tests, for the specification I, are shown in Table 6.[16]

Assessing each of the five equations, the white noise residuals hypothesis is supported by the Q-statistics of Ljung–Box [52], while the ARCH-LM test [24,35] supports the hypothesis that there is no conditional heteroscedasticity and the Jarque-Bera test [7,8,40] supports the hypothesis of residuals with normal distribution.[17] These hypotheses are also supported by the tests in their multivariate versions.

The cointegration vector and loading parameters estimated by three cointegration models are shown in Table 7. It should be observed that the cointegration vector was normalized with respect to the variable crime.

To assess the statistical significance of the deterministic regressors in the cointegration vector, both models specification II e III were reestimated with restriction in matrix $\boldsymbol{\beta}$.[18]

Table 7. Cointegration vector and loading parameters of the VECM(2) with three different specifications.[a]

	Series					Deterministic terms	
Vector	crime	arrest	gun	unemp	wage	Constant	Trend
			Specification I				
$\hat{\beta}'$	1.0	1.3502	1.0784	−1.2854	2.2548		
		(0.1275)	(0.1602)	(0.1466)	(0.3069)		
$\hat{\alpha}'$	−0.279	−0.1475	−0.3936	0.2044	−0.04003		
	(0.1161)	(0.1622)	(0.2058)	(0.05981)	(0.04638)		
			Specification II				
$\hat{\beta}'$	1.0	1.3509	1.0689	−1.3034	2.211	−22.003	
		(0.1275)	(0.1603)	(0.1467)	(0.307)	(3.0683)	
$\hat{\alpha}'$	−0.286	−0.1584	−0.3538	0.2144	−0.02249		
	(0.1155)	(0.1604)	(0.2055)	(0.05948)	(0.04821)		
			Specification III				
$\hat{\beta}'$	1.0	1.593	1.4756	−1.5262	2.0451		0.01046
		(0.1557)	(0.2102)	(0.2093)	(0.4627)		(0.005924)
$\hat{\alpha}'$	−0.2228	−0.1086	−0.3917	0.1765	−0.001278		
	(0.09826)	(0.1367)	(0.1708)	(0.05003)	(0.0393)		

Note: [a]Standard errors in parentheses; series definitions are provided in Table 1.

In specification III, the value test statistic for null hypothesis H_0: $\beta_{\text{trend}} = 0$, with one degree of freedom, is $\chi^2 = 0.99$ (p-value $= 0.32$). Therefore, the hypothesis that there is no linear trend in the long-run model is not rejected. In specification *II*, however, the value test statistic for null hypothesis H_0: $\beta_{\text{cons}} = 0$, with one degree of freedom, is $\chi^2 = 14.41$. In this case, we chose this specification. Therefore, from now on, the analyses refer to the VECM(2) with a constant inside and outside the cointegrating vector.

The next step was that of applying likelihood ratio (LR) tests successively to test the significance of each of the estimated coefficients. The value test statistics, with one degree of freedom, are as follows: $\chi^2_{\text{arrest}} = 21.86$, $\chi^2_{\text{guns}} = 17.91$, $\chi^2_{\text{unemp}} = 14.2$ and $\chi^2_{\text{wage}} = 12.57$. It can be concluded, therefore, that all variables are highly significant in the long-term relationship.

It should be observed that the same signs of long-run relationships in the model estimation with specification II were observed in the two other tested specifications. As suggested by low standard errors in the respective estimates of the coefficients, such relations were also seen to be statistically significant in the LR tests. Therefore, irrespective of the cointegration model with which the long-run relationships were tested, the observed relations are statistically significant. This indicates that the estimations are robust.

Only the crime e unemp variables appear to cointegrate, as the loading coefficients for the other ones have a high standard error. In this context, restrictions should be imposed and tested on the loading coefficient vector. First, we tested the joint null hypothesis that all adjustment coefficients (α) are equal to zero. Under the null hypothesis, with one degree of freedom, the value test statistic is $\chi^2 = 21.97$. A second set of tests was applied for the purpose of testing individual hypotheses. The value test statistics, with one degree of freedom, are as follows: $\chi^2_{\text{crime}} = 4.1$, $\chi^2_{\text{arrest}} = 0.67$, $\chi^2_{\text{guns}} = 1.77$, $\chi^2_{\text{unemp}} = 8.49$ and $\chi^2_{\text{wage}} = 0.18$. It can be concluded that, in fact, only the crime and unemp variables cointegrate. One should consider, however, that the loading coefficient of the variable guns is significant at 18%.

At this point in the analysis, a relevant question emerges that requires further investigation. In the previous subsection, we applied the Perron test to check for a unit root in three series that could have been impacted by the DS. In that subsection, we found a statistically significant structural break in the crime series. It is thus possible that this series contains nonlinearities.

A question arises when the possibility of nonlinearity in the data is plausible: is the nonlinear specification higher than the linear model? The statistical equivalent to this question is: can the linearity hypothesis be rejected in favor of the nonlinear model [37]?

According to Enders [23], any neglected nonlinearity in univariate time series can be checked using the autocorrelation function of the square values of the series, i.e. using the McLeod–Li test [60]. This procedure consists in determining whether there are statistically significant autocorrelations in the square residuals from a linear model. We applied this test to the crime series using the best-fitting linear model, ARMA(p, q), in order to compute the residuals and square residuals. We used the Ljung–Box Q-statistic [52] to determine whether the square residuals had any serial correlation. This statistic is given by $Q = T(T+2)\sum_{k=1}^{n} \rho_k/(T-k)$, where ρ_k is the sample correlation coefficient between square residuals $\hat{e}_t^2$ and $\hat{e}_{t-1}^2$. Value Q has an asymptotic χ^2 distribution with n degrees of freedom if the e_t^2 sequence is nonlinear. Notice that the McLeod–Li test is an exact Lagrange multiplier (LM) test for ARCH errors [24]. However, according to Enders [23], the test has substantial power to detect various forms of nonlinearity.

In the previous subsection, it was observed that the transformed (logarithmic) crime series has a unit root, i.e. it is an integrated of order one – I(1). Therefore, in the presence of a unit root, a plausible set of ARMA(p, q) models was identified using the extended autocorrelation function [85]. In addition to the indicated models, other ones corresponding to several lower-order pairs p and q were estimated and checked. Using the first difference in the crime series, Δcrime_t, the model that best fit the data was an ARMA(2, 2) model, for which the Akaike information criterion [3,4] is -100.10.

The diagnostic tests applied to the residuals of this model were all satisfactory. The hypothesis of white noise residuals is sustained by the Ljung–Box Q-statistics[19] [52] for $k = 6, 8, 10, 12$, since the p-values are 0.50, 0.70, 0.89 and 0.95, respectively; the possibility of residuals with conditional heteroscedasticity is rejected by the ARCH-LM test [24] for $k = 4, 6, 8, 10, 12$, since the p-values are 0.27, 0.29, 0.16, 0.18 and 0.19, respectively; finally, the Shapiro–Wilk test [74,75] sustains the hypothesis of residual normality (p-value $= 0.77$).

It should be recalled that the purpose of this empirical exercise is to capture any possible nonlinearity in the crime series. The p-values of the McLeod–Li test for $k = 1, \dots, 12$ are 0.46, 0.65, 0.57, 0.65, 0.78, 0.63, 0.50, 0.49, 0.59, 0.61, 0.56 and 0.60, respectively. As no Q-statistics is statistically different from zero, the null hypothesis of the McLeod–Li test is not rejected. Thus, there is no potential problem with the linear specification.

Although we did not see any empirical evidence leading us to reject linearity in the crime series, we decided to continue with this investigation by specifically analyzing the threshold autoregressive (TAR) models, in which there are m different regimes ($m > 1$).

Considering that a cointegration relationship was diagnosed previously, the optimal empirical strategy for this investigation would be to perform a test for null hypothesis of linear cointegration against threshold cointegration according to Hansen and Seo [38], but using the five variables of the system as estimated in our study, instead of the bivariate case that they assessed. Unfortunately, all the tests found in the empirical literature are only applied to bivariate systems. This is due to the fact that a grid search is required, whose dimension is a function of the number of variables, making it almost unfeasible for more than two variables to be used.[20] For this reason, in this study we applied two specific linearity tests without taking the cointegration relationship into account.

First, we applied the test suggested by Hansen [37] for the Δcrime_t series and then we applied the multivariate test proposed by Lo and Zivot [53] for the $\Delta\mathbf{y}_t$ series[21]. The former tests linearity against a threshold with bootstrap distribution and the latter test is just the multivariate extension of the first one.

In the first linearity test, an F-test comparing the residual sum of squares (SSR) of each model is computed by $F_{ij} = T((S_i - S_j)/S_j)$, where S_i is the SSR of the model with i regimes (and so

$i-1$ thresholds). Two tests are applied: linear autoregressive (AR) model versus TAR(2) model, and AR model versus TAR(3).[22]

In their respective order, the results of these tests are $F_{12} = 12.82$ and $F_{13} = 26.00$. Making 1000 bootstrap replications, the p-values are 0.32 and 0.42, respectively. Thus, we cannot reject the hypothesis of an AR model against the TAR(2) or TAR(3) models. This suggests that the linear model is the appropriate one for the Δcrime_t series.

As already mentioned above, the second linearity test is nothing but a multivariate version of the first one. Instead of an F-test comparing the SSR for the univariate case, an LR test comparing the covariance matrix of each model is computed by $\text{LR}_{ij} = T\{\ln(\det \hat{\mathbf{\Sigma}}_i) - \ln(\det \hat{\mathbf{\Sigma}}_j)\}$, where $\hat{\mathbf{\Sigma}}_i$ is the covariance matrix estimated by the model with i regimes (and so $i-1$ thresholds). The same possibilities of the univariate test were evaluated.

The results of these tests are $\text{LR}_{12} = 37.93$ and $\text{LR}_{13} = 115.86$, respectively. Once again, we made 1000 bootstrap replications and obtained p-values equal to 0.89 and 0.25, respectively. Thus, we cannot reject the hypothesis of linearity in favor of a nonlinear model either. Therefore, the results indicate that the linear specification is superior to a nonlinear model for modeling $\Delta\mathbf{y}_t$ series. We will report and discuss the estimates of the linear cointegration model in the next section.

5. Results and discussions

Table 8 reports the estimates obtained for the equation Δcrime_t by the model identified in the previous section, i.e. by VECM(2) with the specification II.

First, it is worthy of mention that the estimated coefficient of the dummy intervention variable for the DS is negative (−0.129) and significant at 1% ($t = -3.096$), indicating that disarming the population led to a reduction in the lethal crime rate. The statistical significance of this dummy variable corroborates the evidence provided by Santos and Kassouf [77] that disarming citizens pushed the lethal crime rate down. If this dummy variable had been significant in the other equations as well (i.e. for the other system variables), there would be signs of a spurious relationship, particularly if it were significant for the economic variables, with which the disarmament policy should not have any relationship. At conventional levels, however, apart from significant in the crime rate equation, this relationship was only meaningful in the arm series equation, which is plausible. After the law on disarmament was regulated, there are less seizable firearms in circulation, either because fewer people are carrying guns on the streets or because firearms were voluntarily turned in by citizens in large numbers during the disarmament campaign.[23] Therefore, even considering other possible causes of crime drop in São Paulo in the model specification, the hypothesis that the DS contributed to it is not rejected.

Regarding the short-run relationship, evidence shows that economic conditions and police performance have no statistically significant effects on the lethal crime rate. In terms of the variables tested in this study, the absence of short-run effects is plausible. There is the possibility that a reasonable time interval might be required for economic changes to influence decisions of new individuals to engage in crime or not.

Cantor and Land [10] discussed this issue and stated that changes in economic conditions are unlikely to have short-run effects on crime, especially because in most cases downturns in economic activitiy are compensated by government actions. In Brazil, for example, an unemployed individual in the formal market is entitled to welfare (unemployment insurance) for a certain period of time (several months) that is considered sufficient for him or her to find a new job. Moreover, in most cases, financial support is provided by family, friends or charities. Therefore, according to Cantor and Land [10], it is impossible to say that economic recession immediately motivates people to commit criminal offenses. However, according to them, an unemployed person might

Table 8. Estimates of short-run relationship[a] Resposta: Δcrime_t.

Variable	Coefficient	Standard errors
α	−0.286**	0.1155
DS	−0.1286*	0.0415
S_1	0.00306	0.05457
S_2	−0.06438	0.0607
S_3	−0.1014***	0.0553
Δcrime_{t-1}	−0.08359	0.1713
Δcrime_{t-2}	0.02545	0.1658
Δwage_{t-1}	0.005803	0.437
Δwage_{t-2}	0.1376	0.3849
Δunemp_{t-1}	0.3269	0.2871
Δunemp_{t-2}	0.13066	0.3192
Δguns_{t-1}	0.1168	0.1337
Δguns_{t-2}	0.05866	0.1016
$\Delta\text{arrest}_{t-1}$	0.2110	0.1644
$\Delta\text{arrest}_{t-2}$	0.2054	0.141
R^2	0.4527	
Observations (n)	53	

Notes: [a]The standard errors of $\hat{\alpha}$ was adjusted to degrees of freedom.
*Significance at 1% in t statistics.
**Significance at 5% in t statistics.
***Significance at 10% in t statistics.

become more willing to engage in crime than an employed one when social security benefits and other sources of financial and psychological support have been exhausted. In this context, it is possible that economic status has noshort-run effects on the crime rate, but rather long-run ones. Likewise, economic conditions might require a certain time interval to actually influence the opportunity cost of crime. This cost, according to Becker [6], is measured by an individual when deciding to commit a crime or not.

Non-statistically significant relations between economic conditions and crime have been pointed out and discussed in previous studies. Scorcu and Celline [81], for example, observed that the growth rate of economic variables does not affect the short-run dynamics of crime rates significantly. According to those authors, the main effect of economic conditions on crime seems to arise from adjustment processes, which lead current crime rates towards their long-run equilibrium level. In general terms, the authors concluded that the crime rate is not correlated to the economic variables considered in their study. Based on such evidence, they suggested that the attention paid in most studies to the short-run effect between economic activity and crime is perhaps excessive and that more attention should be paid to long-run relationships. Saridakis [79] also addressed the issue of the short-run effects of economic status on crime. Using data from England and Wales, the author concluded that, of all economic variables analyzed, only the unemployment rate was seen to have statistically significant short-run effects on violent crime rates. However, as stressed by the author, the effect is numerically very low, suggesting that violent crime rates do not adjust to economic conditions in the short run. Also, as suggested by Field [27], the short-run determinants of crime can differ from long-run determinants, making it more difficult to identify the causes of crime.

Since short-run relationships are not significant from the statistical point of view, the focus of this study will once again be placed on the long-run determinants of crime. With the aim of analyzing the long-run relations between crime and other variables in the model, vector $\hat{\boldsymbol{\beta}}'$ of the model selected in the previous section was rewritten in the form of a relation determining the crime rate:

$$\widehat{\text{crime}}_t = 22.003 + 1.3034\text{unemp}_t - 2.211\text{wage}_t - 1.3509\text{arrest}_t - 1.0689\text{guns}_t. \quad (5)$$

Owing to methodological differences in this study, its results are not directly comparable to those of other studies that investigated the causes of crime in Brazil. Nevertheless, we intend to discuss them, as much as possible, in the light of economic theory, based on which the first studies on the topic were conducted [6,21], and of prior domestic and international empiric evidence.

Firstly, it is worth saying that the signs of all the coefficients of the Equation (5) correspond to the relationships predicted in these economic approaches to criminal behaviors, although what is being dealt with is not the original crime supply equation proposed by Becker [6] and extended by Ehrlich [21].

The crime rate was shown to be positively related to unemployment and negatively related to wages. It is worth remembering that these two factors are proxy variables for both for the opportunity cost of crime and for expected return on crime. As for the proxy variables for police enforcement, it was shown that increases in the relative number of arrests of drug traffickers and in the relative number of seized firearms led the crime rate to drop.

Unemployment and wages are often seen, especially by economists, as important determinants of criminal behavior. In this context, hypothetically, improvements in labor market conditions in the state capital explain, at least partially, the reduction in the crime rate in São Paulo.

The effects of labor market conditions on crime levels have motivated empirical economic studies even before the classic theoretical study by Becker [6] was published. Before that author, Fleisher [28] produced the first empirical evidence on the relationship between labor market conditions and crime.

As far as we know, there are no specific empirical analyses of the above-mentioned relationship in the Brazilian literature. However, it has been a common approach to use a measure of market conditions to control for the opportunity cost of crime in estimating equations for crime ‘supply’ (see [5,26,33,63,76], and others).

According to economic theory on the causes of crime, better conditions in the labor market increase the opportunity costs of crime. This in turn reduces the probability of an individual committing a crime. As shown in Figure 1, unemployment levels in São Paulo have been trending downward approximately since the twenty-eighth quarter of the sampling period. As shown in the same figure, this drop took place concomitantly with an upward trend in real wages, which became more pronounced at about the thirtieth quarter.

The evidence that a lower unemployment rate reduces crime levels is in tune with the findings of most previous studies on this relationship ([79,81], and others), but not with those of authors that did not find a significant relationship between them [34,51,64].

With regard to real wages, the result is consistent with the evidence produced by Gould *et al.* [31] when they analyzed the effects of labor market opportunities on the crime rate in the USA. It is also in tune with the results of a study by Santos and Kassouf [76], although they used a very different measurement to reflect labor market conditions. Using a measure of job turnover, it was concluded that a heated labor market pushes crime levels down.

It is important to consider that the effects of the economic variables analyzed here might have a major bearing on the decision of individuals who could potentially become criminals in the future and of individuals serving sentences who will have to decide if they will work or relapse into crime after being released. But it is very likely that conditions prevailing in the labor market are not considered in the decision of habitual criminals to commit a crime. In other words, better conditions in the labor market are probably more effective to prevent new individuals from engaging in criminal activities than to induce active criminals to give up their criminal 'career'. For those already engaged in crime, the deterrent effects of the public security policy are more effective in reducing the number of crimes committed.

Another hypothesis put forward particularly by public security policymakers is that a better performance of law enforcement is the main cause of the crime drop observed in São Paulo.

According to the classic economic theory of crime of Becker [6], society's objective is that of minimizing damages caused by crime by inducing individuals to commit crimes at an 'optimal' level. It does so through its legal representatives, i.e, policymakers. These, in turn, can choose the amount of funds to be allocated to public security and how. This decision directly affects the probability of failure of a criminal act. The legislator is also the one who defines the forms and severity of penalties applied to convicted criminals. In this context, it is plausible to assume that both the variables arrest and guns reflect these choices indirectly, since they were built from indicators of results of law enfocement activities.

In this context, the evidence that shocks in the relative number of arrests for drug trafficking are negatively transmitted to the crime rate is plausible. Referring to Figure 1 once again, note that this series, which trended downward until about the fifteenth quarter of the sample, took a sharp upturn while the crime rate began to fall. It can thus be inferred that fighting drug trafficking also leads to a decrease in other crimes, particularly lethal crimes. It should be recalled that this variable was included in the empirical model specification for two reasons: on the one hand, it reflects the results of police activities, particularly in fighting organized crime; on the other hand, it is a proxy for the presence of illegal profitable activities, which are usually managed and maintained through violence and corruption.

Based on the estimation results, the hypothesis that an effective fight against drug trafficking is co-responsible for the significant drop observed in the crime rate in São Paulo is not rejected.

There is a heated discussion on the role played by incarceration of low-danger criminals in the dynamics of crime, but when it comes to drug dealers, incarceration is consensually seen as the only effective penalty for reducing crime levels. Disarming the population, particularly criminals, is the key for reducing crime.

The arrest series reflects the relative amount of arrests for drug trafficking. In the drug market, arrests of drug dealers, in particular of leaders of drug gangs, imply arrests of other individuals who are active in criminal organizations. Therefore, it is plausible that shocks in this variable are permanent. One must also consider that when a criminal is arrested, there will be less people commiting crimes. Therefore, unless new individuals engage in a criminal activity or the 'technology' used in that activity evolves, fewer crimes are expected to occur after an arrest.

The hypothesis that illegal activities also specialize over time is plausible. Learning by doing in criminality activity reduces the costs of planning and committing a crime, apart from reducing the probability of failure in each crime (likelihood of a crime being registered and of the individuals involved being arrested, tried, convicted and punished). Consequently, it increases the expected return on criminal activity. Moreover, the longer an individual is engaged in illegal activities, the lower the expected returns on lawful activities, and thus the lower the opportunity cost of crime for him or her. One must also consider that the probability of crime failure in Brazil is undeniably low. A strong sense of impunity further increases the expected return on crime. Together, all of these factors increase the likelihood of recidivism and the probability of an individual becoming a habitual criminal. Therefore, a rise in the rate of arrests brings about an important dissuasive effect on the decisions of new individual to engage in a criminal activity or not.

Regarding the second measure used in an attempt to capture the effect of police performance on crime, i.e. guns, it is considered that the increase in the relative number of firearms seized cause negative effects on the crime rate, which is in tune with the findings of most studies that investigated a possible relationship between the availability of firearms and the crime rate. In this literature, most authors support the 'less guns, less crime' thesis ([19], and others), which is, however, rejected by others ([54–57], and others).[24] It should also be stressed that the evidence presented in this study corroborates the conclusions of Cerqueira [12].

One must consider that the guns series reflects two types of results. On one hand, it shows that the more firearms are seized in a given period, the lesser illegal firearms will likely be seized in the future. On the other hand, it suggests that the amount of firearms, a factor closely related to the number of searches and identification of persons by police officers, reflects police performance in disarming people carrying firearms illegally and in disarming criminals. Therefore, the statement that this variable causes negative effects on the crime rate suggests that disarmament is necessary for reducing the crime rate in other locations, as was done in the case of São Paulo.

6. Concluding remarks

This study provides evidence of the long-run effects of crime determinants. In particular, it pointed out the major role played by unemployment in explaining crime levels. Its main contribution, however, was that of shedding light on the possible causes of the sharp drop observed in the crime rate in São Paulo. We believe that it was a major step in this investigation. Further investigations are essential, particularly for validating the results presented here. Another contribution of this study to the empirical literature is the fact that it generated more knowledge about the non-stationarity and cointegration properties of the crime rate. The results of the study were discussed in the previous section. However, it's worthwhile summarizing those seen as the most relevant ones.

The first one is that the DS continued to have a significant negative effect even after other alleged causes of crime reduction in São Paulo were considered. This result reinforces the previous evidence in Santos and Kassouf [77].

There were no observations of statistically significant short-run effects on the lethal crime rate, the unemployment rate, real wages, and proxies for police performance, but there is statistical evidence that these variables have long-run effects on criminality. The long-run relationship identified between economic conditions and the crime rate indicates that lower unemployment rates and increases in real wages were determinants of crime reduction in São Paulo. Effectively fighting drug trafficking by arresting drug dealers and seizing firearms were as important as ensuring more favorable economic conditions to the population.

By studying the clearly successful experience of São Paulo in reducing crime, it became clear that economic activity is closely related to criminal activity levels in the long run. Extrapolated to the country, this piece of evidence suggests that apart from enhancing deterrence of criminal behavior, thus increasing the likelihood of crime failure, the Brazilian government should adopt policies to improve the Brazilian labor market, especially by taking effective measures to reduce unemployment even more. This will increase the opportunity cost of crime, thus discouraging new individuals from engaging in crime and maybe reducing the recidivism rate.

In the previous section, it was pointed out that labor market conditions are not very likely to be taken into account by hardened criminals before committing a new crime. It is not very plausible that a reduction in unemployment and higher earnings from legal activities can cause criminals, that is, people already engaged in criminal activities, to stop committing crimes. It should also be considered that criminals do not suffer the direct effect of policies designed to disarm the population. They only feel it indirectly, i.e. with fewer weapons in circulation, they are less likely to be stolen or robbed. Therefore, to reduce the high crime rates in most of the Brazilian territory, effective public security policies are fundamental. In this context, increasing the rate of imprisonment for crimes committed is essential for removing criminals from the streets and discouraging individuals who are likely to participate in criminal activities from engaging in them. The first step, however, is reviewing the Brazilian prison system, the severity of penalties, and the quality of public security institutions. If this is not done, more arrests in the present will not necessarily lead to less crimes being committed in the future.

7. Computational details

All empirical procedures were carried out using the **R** statistical package [73]. The Lowess smoothing of the series was done in the *stats* package [73]; the Box–Cox analyzes were performed with the *FitAR* package [61]; the Dickey–Pantula test equations and the Perron test were written and estimated using the *dynlm* package [87]; both the ADF–GLS and the KPSS tests were carried out using the *UnitrootUrcaInterface* package [86]; the likelihood ratio tests on restrictions in the system were carried out in the *urca* package; the cointegration analyses and diagnostic tests were carried out using the functionalities available in the *vars* package [69]; the estimation of ARIMA models and the Shapiro–Wilk test were all carried out using the *stats* package [73]; the *TSA* package [13] was used to calculate the extended autocorrelation function; the ARCH-LM and McLeod–Li tests were performed in ArchTest [32] and *TSA* packages [13], respectively.

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Notes

1. Federal Law 10,826 of December 2003, known as the DS, which was regulated in July 2004, disciplined the possession and carrying of firearms in Brazil, provided for stricter penalties for illegal possession and carrying of firearms, and made it possible for disarmament campaigns to be carried out. Its full text is available at: http://www.planalto.gov.br/ccivil_03/Leis/2003/L10.826.htm.
2. 2004Q1 denotes the first quarter of 2004.
3. For a complete explanation of the econometric methodology that was applied, see [23,35,46], among others.
4. Actually, the lethal crime rate and the number of firearms seized by one hundred thousand population followed very similar long-run paths in the sampling period under analysis. The Pearson correlation coefficient estimated between them was 0.96 ($t = 24.42$).
5. Crimes that include many behaviors related to the sale or distribution of substances that can cause physical or psychological addiction.
6. Data available at http://www.ssp.sp.gov.br/esttistica/trimestrais.aspx. Unfortunately, these data are published separately by quarter in pdf file. We would like to acknowledge that the data for the number of searches and identification of persons in periods preceding the first quarter of 2005 were kindly provided by Dr Tulio Kahn, former coordinator of the Coordinating Office for Planning and Analyses of the Public Security Department of the State of São Paulo. For details on the process of collecting and interpreting the data, see [11].
7. Saridakis [78] uses the unemployment rate among men. Besides this measure, we considered the possibility of using the unemployment rate only among family heads. However, no significant difference was observed in the time path of these three series.
8. Data available at http://www.seade.gov.br/produtos/ped/.
9. The minimum wage is deflated by the cost-of-living index in São Paulo, which is calculated by DIEESE. Other details can be found at http://www.dieese.org.br/esp/metodsm.xml.
10. We also applied the statistical criterion suggested by McLeod [62] to monitor the variance behavior in successive differences in the series. Simulations by the author suggest that an excessive number of differences result in a negative first-order autocorrelation value in the differentiated series tending to -0.5; when, however, the series is correctly differentiated, variance in the transformed series decreases; on the other hand, excessive differences increase variance.
11. The concomitant use of both tests is justified by the fact that, by reversing the null hypothesis, the first test reduces the low-power problem of the second, particularly as the autoregressive coefficient is closest to one.
12. To our knowledge, in Brazilian literature only Pereira and Carrera-Fernadez [66] had made a similar attempt, but using the methodology proposed by Engle and Granger [25]. The authors reached the conclusion that there is a cointegrating relationship between crime (total crime and vehicle theft or vehicle robbery) and income inequality.

13. Note that the constant outside the cointegrating vectors represents a linear trend at the level of $\mathbf{y}_t$. For details on five possible Johansen's specifications, see Johansen [44,46] and Harris [39], among others. Johansen [45] discusses the role of the deterministic regressors in a cointegration relantionship.
14. It should be observed that it is not different from the maximal eigenvalue test.
15. See Johansen *et al.* [48] for technical details as to how the two models are defined.
16. It is noteworthy that in the other two specifications, I and III, the diagnostic test results do not differ from the results obtained for the specifications II.
17. A multivariate version of this test can be computed by using the residuals that are standardized by a Choleski decomposition of the variance–covariance matrix for the centered residuals. In this case the test result is dependent upon the ordering of the variables. We used the ordering reported in the next section. To check the result, this test was redone with different sorts of model variables. There was no difference between the test results regarding the conclusion.
18. For details about restriction tests on $\boldsymbol{\alpha}$ and $\boldsymbol{\beta}$ see Johansen and Juselius [47] and Johansen [46], among others.
19. Statistics with reduction of four degrees of freedom derived from order $p + q$ of the model.
20. See Stligler [84] for an overview of threshold cointegration.
21. The variable definitions are provided in Table 1.
22. It should be noted that the value of the threshold, (τ), is unknown in these tests. It is estimated along with other parameters of the TAR model. In others worlds, these are 'endogenous nonlinearity' tests, in the sense that the transition variable is a function of the system variables. Following Hansen [37], we used $\tau = 0.1$ for the minimal percentage of observations in each assumed regime and, based on the linear model specification used in the previous linearity test, we used two lags in the specification, i.e. $p = 2$.
23. According to the Ministry of Justice, national campaigns have, since 2004, removed around 570,000 firearms from circulation. In its 2011 edition, nearly 34,000 firearms were withdrawn from circulation, approximately 29.3% of which were in the state of São Paulo. Information available at http://portal.mj.gov.br; accessed on 01/13/2012.
24. McDowall [59] brings together and discusses the main points of this controversial discussion.

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