TO FINE OR TO PUNISH IN THE LATE MIDDLE AGES

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To Fine or to Punish in the Late Middle Ages
A Time Series Analysis of Justice Administration in Nivelles, 1424-1536

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Abstract

We use recently developed time series techniques to analyse long historical data on the budget of the Court of Nivelles (Belgium) in the late Middle Ages. The unidirectional causality from receipts to expenditures supports the idea that the medieval system under consideration is a fine-focused (tax-oriented) control system. However, empirical evidence on the presence of changes in receipt collection around 1520 favours the hypothesis of the decline of the medieval, urban-based, tax-oriented system and its transformation into a punishment-focused system.

JEL Classification numbers: C52, H72, K14, N43

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Issues

The Middle Ages are frequently considered as a period of obscurantism and used to be well known as the "Dark Ages". Actually, the medieval institutions were more sophisticated than suggested by this simplistic representation. One impressive administrative output of this period is the bulk of domanial accounts of Kings and Landlords, especially in England but also in the Netherlands or Low Countries (the present Benelux).

The accounts of justice officers were part of this domanial accountancy. They reveal a full-fledged justice "budget" with receipts and expenditures. For the Netherlands, perhaps 100,000 of such periodic accounts are available, spanning the period between the middle of the 13th century and the middle of 17th century for Artois, Flanders, Brabant, Hainaut, Namur, Luxemburg, Holland, Zeeland, Gheidses, etc. The accounts of local, regional and principal officers provide us with a substantial information about the financing of the administration of justice.

Rousseaux (1990e) analyses various data series on the costs and benefits of justice in the late Middle Ages and Early Modern Times as well as series on prosecution and sentences for 1500-1650. He attempts to rebuild the structural evolution of "punitive" systems on the basis of a local case. For the city of Nivelles, he observes three steps in this evolution: First, the construction of a local system of tax and fine collection during the period 1423-1457. Second, the stabilisation and effective functioning of this taxation system from 1457 to 1520. Third, a crisis in the system and its progressive collapse between 1520 and 1536. From these observations, based on the accounts of the officer of Nivelles, Rousseaux (1990e) proposes a general hypothesis: the construction of a "tax-oriented" system in a seigniorial city starting in the first decade of the 15th century, and its progressive evolution towards a more "punishment-oriented" system.

The aim of this paper is to provide some empirical elements on (1) the existence of a medieval "tax-oriented" system in Nivelles and (2) its progressive change towards a more "punishment-oriented" system around 1520-1530.

Point (1) implies that the main aim of this judicial system is to apply fines quickly and frequently as an immediate sanction for unauthorized behaviour. It is a profitable system which provides receipts to Landlords and cities. The expenditures are generally linked with the collection activity (sergeant's wage, ...) and sometimes include prison wardens expenditures. If we analyse the budget data in a time series perspective, Point 1 implies the existence of unidirectional causality from receipts to expenditures. If receipts and expenditures turn out to be non-stationary processes, causality analysis has to be carried out within a framework allowing for cointegration.

Point (2) is to be related to the emergence of a modern, State-based, judicial system. This system makes a large use of physical sanctions and focuses its action on a smaller group of people. For obvious reasons, it is less profitable than the previous one. Note that this second assumption is quite new in the literature since very few studies on criminal history ¹ have focused on this period (exceptions are Muchembled (1992) and Chiolfoue (1974)). Those who have centered their interest on this period, e.g. as Brouwers (1965) and Marounék (1986), do not succeed in proposing a precise chronology of the mutation.

¹General overviews are in Cockburn (1977), Brewer and Styles (1980), Soman (1980), Gatrell et al. (1980); see also Dupont-Bouchat (1976) for Belgium and Weisser (1982) for a broad synthesis.
This mutation took place in a deep changing context in European societies at the end of the fifteenth and the beginning of the sixteenth century. From an economic point of view, the general situation went bad at the end of the fifteenth century. The demographic decline, the fall in agricultural production and the decline of prices and salaries point to a fall in Europe's growth. A rapid but short recovery occurred during the years 1520-1550 in the Netherlands followed by a structural crisis due to socio-religious and political factors. Various studies (Van Der Wee (1963) ...) confirm this decline for the Netherlands, especially for Brabant. Unfortunately, we do not find in the archives any evidence of such a crisis for Nivelles, but it seems likely that the city's economic prosperity of the 14-15th centuries, based on textile and cloth industries, was challenged in the years 1480, confirming the impact of the general trend on the local market.

From a socio-religious point of view, Christian belief played an important role of "social cement" in the Western community. The Protestant Reformation led to a crisis in social relationships, especially in the Netherlands. Lutherans, and later Anabaptists and Calvinists, were prosecuted by the kings of Spain Charles the Fifth and Philip the Second from 1520 up to the Dutch Revolt and the divorce between Southern (Spanish) and Northern Netherlands (United Provinces) in 1566-1585. This religious dissent had an economic background. Craftsmen and capitalists, especially in the cities, were the most important supporters of the Reformation (Van Der Wee (1963)).

From a political point of view, the beginning of the sixteenth century was characterised by an increase in the interventionism of central political authorities (kings and princes) in local government. This trend was particularly clear in the administration of criminal justice. As a response to the Lutheran Reformation, Charles the Fifth signed a large number of edicts: the "placards" against the heretics. This was a great innovation: For the first time, a general criminal legislation was promulgated in all the Low countries (see Rousseaux (1990a) and (1990b) and Muchembled (1992)). Part of this legislation appeared in the local records of the judges of Nivelles. At the same time, legists like Philippe Wielant and J. de Damhouder (1554) in Flanders were developing modern ideas about the judicial system (public action, inquisitorial written procedure, torture, physical punishment) based on Roman law. Universities started to have a prominent role in the diffusion of Roman law (Van Caenegem (1981))). This seizure of justice by the central State culminated during the Troubles, when in 1570 Philip the Second edicted his "Criminal ordinances" ruling the administration of criminal justice. The 16th century saw the simultaneous development of repressive legislation, increasingly sensitive authorities and perhaps an "acculturation" on the part of the people to religious and moral deviancy (religious dissidents, witches and members of fringe groups were hunted) as well as the birth of a "punitive model" based on corporal punishment and public disgrace. "The advancement of the movement to prohibit feuds, forbidding all forms of revenge, including the feud and the consolidation of territorial states assuming sole power over life and death determined this development" (Van Dülmen (1990), p133). This evidence reinforces the analysis of the interaction between the mode of punishment and the economic environment proposed by Rusche and Kirchheimer (1939) in a Marxist perspective. In this paper their hypothesis can be related to the shift, between 15th and the 16th century, from justice by taxation (fines) to punitive justice.

Beside these political, religious and economic explanations we want to test econometrically the potential occurrence of the structural break implied by Point (2). At this stage,
the exact nature of the structural break is unknown: we do not know if the mutation of
the system should affect all possible relations between the variables or whether it will be
still possible to isolate some invariant relationship. To go beyond the simple search for
structural break, we use the intimate relationship between structural invariance and (su-
per)exogeneity proposed by Engle and Hendry (1993). Given the potential non-stationary
property of the series, the implications of cointegration should be taken into account.

The paper is organized as follows. In the first section, we present the data sources and
the basic functioning of the budget of the Court of Nivelles. In Section 2, the statistical
properties of the data are analysed. The long-run relationships between receipts and
expenditures and their causality links are investigated in section 3. In Section 4 we
address the problem of structural stability of the model and its link with the collection of
fines. Section 5 concludes.

1 Data Sources

We investigate the case of the justice officer of a seigniorial jurisdiction (Nivelles, Walloon
Brabant) of 4,000 – 6,000 inhabitants during the Middle Ages and the Ancient Regime
(14th - 18th century). A local seigniorial officer supervised the functioning of the Court on
behalf of the four powers involved in the local administration of justice: The municipality
and the three landlords (Duke of Brabant, Abbess of St. Gertrud, Provost of St. Gertrud’s
abbey). The officer controlled the receipts and expenditures, and the destination of final
receipts (or expenditures) was controlled by the four authorities.

Figure 1: The Court Budget

The way the various receipts and expenditures are shared among the authorities is
presented in Figure 1. Common receipts (fines) and Common Expenditures (preventive
imprisonment and management of justice) are locally allocated between the Municipality
and the Landlords.

The sources are conserved in the Belgian State Archives (Brussels) and are made up of
128 periodic accounts from 1423 to 1550. These accounts are complete for the period 1423-
1536, i.e. for more than one hundred years. We faced three problems in homogenising
the series:
1. *The diversity of the structure of the accounts*: The internal structure of this accountancy changed during this long period. The structure is unstable from 1423 to 1457. In 1457, the general structure is fixed and remained more or less the same up to 1536. To avoid the problems linked with the definition of the various items, we limit our analysis to aggregate series: number of fines (Figure 2) and common receipts and expenditures (Figure 3).

3. *The diversity of the periodicity of the accounts*: Most of the accounts cover one year, but some covered a few weeks and others six years. We homogenise these series using a three period moving average.

4. *The monetary diversity* is problematic. However, each account was summarized in one monetary unit. For the first years (1423-1448), this unit was the “pound” (“1 lb = 20 sous = 240 deniers”). After 1448, the monetary unit was the Brabants “plaque” (“1 plaque = 6 gros = 576 deniers = 24 mites”). In the 16th century a new monetary unit appeared: the “florin carolus” created by Charles V. (“1 fl = 20 patars = 240 deniers”). The initial relation between these currencies is presented in Table 1.

<table>
<thead>
<tr>
<th></th>
<th>plaques</th>
<th>pound</th>
<th>florin</th>
</tr>
</thead>
<tbody>
<tr>
<td>livre</td>
<td>0.83</td>
<td>1</td>
<td>0.28</td>
</tr>
<tr>
<td>plaques</td>
<td>1</td>
<td>1.20</td>
<td>0.33</td>
</tr>
<tr>
<td>florin</td>
<td>3</td>
<td>3.60</td>
<td>1</td>
</tr>
</tbody>
</table>

As a consequence of the fluctuating value of monetary units, these relations changed over time. For the analysis, we expressed all data in plaques. From 1438 to 1536, this unit of account seems to remain stable.

Figure 2: Number of Fines
2 Statistical Properties of the Data

As it is clear from Figures 2, 3 and 4, our data series display trending behaviour. Consequently, before searching for stable relationships between our series, we have to test whether or not they are trend stationary or integrated processes. This is a very important preliminary since standard asymptotic distribution theory often does not apply to regressions involving stochastic non-stationary variables, and statistical inference can be misleading if non-stationarity is ignored. In general, regressions involving non-stationary variables are “spurious” (Granger and Newbold (1974)) with a very low Durbin-Watson statistics and a high $R^2$ giving the false impression of a good fit. In this case regression coefficients converge to random variables (instead of converging to the true coefficients) as the sample size increases to infinity (Phillips (1986), see also the discussion in Campbell and Perron (1991)).

In order to test for the presence of stochastic nonstationarity in our data, we first investigate the integration order of the series using two different unit root test: the Augmented Dickey-Fuller (ADF) tests and Phillips-Perron’s (1988) non-parametrically modified Dickey-Fuller test. The ADF test is based on the following regression model:

$$y_t = \alpha + \mu t + \rho y_{t-1} + \sum_{i=1}^{p} \psi_i \Delta y_{t-i} + \varepsilon_t$$  

(1)

This regression model assumes that the univariate time series $y_t$ can be approximated by a finite order autoregressive model (of order $p + 1$). If $p = 1$, the shocks $\varepsilon_t$ have a permanent effect on $y_t$. Several assumptions concerning the deterministic components can also be tested using one-sided (F-type) tests whose critical values are given in Fuller (1976). The limiting distribution crucially depends on the deterministic components of the univariate series.

Phillips (1987) non-parametric correction to the Dickey and Fuller test (based on (1)...
without lagged first differences) is motivated by the fact that the ADF tests are only valid under the crucial assumption of i.i.d. processes for $\varepsilon_t$. In practice, it may be more realistic to allow for some dependence among the $\varepsilon_t$'s (e.g. allowing for serial correlation and heteroskedasticity but also controlling the extend of temporal dependence) so that, although there may be substantial dependence among recent events, events which are separated by long intervals of time are almost surely independent. Phillips and Perron (1988) have extended Phillips approach to the case of non zero drift and deterministic trend while Ouliaris et al. (1988) allow even for higher order trend polynomials.  

From the figures reported above, it appears clear that most of our series are likely to possess some deterministic trend component or might even be characterised as trend stationary processes. Therefore, we implicitly followed the sequential testing procedure proposed by Perron (1988) which aims to start from a quite general specification with both trend and constant term. The first hypothesis which is tested is that of a random walk (unit root) with drift against a trend stationarity stationary process. In the case of non-rejection, one then tests for the significance of the trend term and so on. The final hypothesis to be tested in this sequence (if all the previous less restrictive hypotheses have not been rejected) is the driftless random walk against the simple zero-mean stationary AR(1) process. The results of the unit root tests are reported in Table 2. For both ADF-type and Phillips-Perron type statistics, the lag truncation was set at four after having scrutinised the autocorrelation function of the first differenced processes. The following notation is used: $\tau$ is the usual t-test based on (1) without trend term, $\tau_\mu$ is the t-test based on (1) with trend term. The $\Phi_i$, $i = 1, 2, 3$ are one sided F-type statistics respectively testing the null hypotheses of (a) $(\mu, \rho) = (0,1)$ in (1) without trend, (b) $(\mu, \alpha, \rho) = (0, 0, 1)$ in (1) with trend, and (c) $(\mu, \alpha, \rho) = (\mu, 0, 1)$ in (1) with trend. The "Z" versions reported in the Table have the same interpretation but are computed with the Phillips-Perron non parametric correction instead of the lag augmentation.2

The variables (expressed in logarithms) are denoted $f$ for the fine number and $r$ and $e$ for common receipts and expenditures.

<table>
<thead>
<tr>
<th>Test Statistic</th>
<th>$\tau$</th>
<th>$\tau_\mu$</th>
<th>$\Phi_1$</th>
<th>$\Phi_2$</th>
<th>$\Phi_3$</th>
<th>$Z(\tau)$</th>
<th>$Z(\tau_\mu)$</th>
<th>$Z(\Phi_1)$</th>
<th>$Z(\Phi_2)$</th>
<th>$Z(\Phi_3)$</th>
<th>95% critical value</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>-2.73</td>
<td>-3.11</td>
<td>3.74</td>
<td>3.26</td>
<td>4.88</td>
<td>-3.36</td>
<td>-4.70</td>
<td>5.65</td>
<td>4.05</td>
<td>6.07</td>
<td>-3.45</td>
</tr>
<tr>
<td></td>
<td>-1.94</td>
<td>-3.96</td>
<td>2.04</td>
<td>5.35</td>
<td>7.87</td>
<td>-3.27</td>
<td>-5.74</td>
<td>5.36</td>
<td>11.02</td>
<td>16.50</td>
<td>-4.71</td>
</tr>
<tr>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
</tbody>
</table>

As shown in this table we reject the null of a unit root with drift against the trend

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2 The long-run variance used in the non-parametric correction is obtained using the Newey and West (1987) non-negative estimator with a truncation lag of 4.
stationary alternative hypothesis for $f$. The same battery of tests applied to the two remaining series confirms, at least at this stage, that our data set contains two I(1) (possibly with some deterministic terms) variables ($r, e$). One should also note that according to the Phillips-Perron test, these two series might almost equally well be characterised as I(0) processes. This phenomena could eventually be explained by various factors including how the data were contracted. However, the point estimates of the first order autoregressive coefficient are, for $r$ and $e$ near unity. We shall therefore follow the advise of, a. o., Campbell and Perron (1991) and consider $r$ and $e$ as integrated processes of order 1.

Any attempt to build conditional or simultaneous equation models linking these different variables should thus clearly recognise the individual properties of the data. In particular, models linking common receipts and expenditures should be built by means of cointegration techniques to explicitly take the stochastic nature of their trend behaviour into account. An alternative method using only univariate statistics could be to reexamine the series of $e$ and $r$ with segmented trends for the period 1423/57, 1457-1520, 1520/36. Indeed, the fact that $r$ and $e$ are I(1) can be related to the two known structural breaks in the system concern the periods 1423/1438 and 1520/1536. The historical context can explain these two changes. In 1438, a normative text from the Duke of Brabant has increased the amount of the fines in Nivelles by 100%. Before 1438 one fine of 1 obole (real unit money) was evaluated to 21 plaques 2 gros. After this date, one obole became evaluated to 42 pl. 4 gros. The break is also due to a structural reorganisation of judicial taxation. In the years 1520, the receipts tend to decline faster in contrast with the appearance of physical punishment, registrated not in receipt but in expenditure, given the costs of physical executions (prison, torture, hanging, etc. ...) However, the theoretical developments of segmented trend functions with multiple breaks is yet not well developed so that we leave this for further research.

3 Long-Run Relationships

In this section, we focus on receipts and expenditures of the common budget, and we investigate if, although these variables are non-stationarity, there exist a linear combination of them which is stationary, and which can therefore be analysed with standard methods. (Trehan and Walsh (1988) apply a similar methodology on US government’s budget data for the period 1890-1986.) In other words, we are interested in finding long-run relationships that are not affected by structural breaks even if each variable taken alone is affected by them. If there exists such linear combination, this one can be interpreted as a long-run relationship that holds between our variables over more than one century, in the sense that the their non-stationarity are compensating. Note that at least one variable will have to adjust in order to satisfy this relation in the long run.

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3 which pointed out that it may be preferable to assimilate near-I(1) series to I(1) ones as their asymptotic behaviour is more adequately described by that of unit root processes than by that of stationary processes.
Multivariate Analysis

Using the methodology of Johansen (1988) and Johansen and Juselius (1990), we estimate a system composed of two autoregressive error correction equations. The determination of the number of stationary linear combinations entering this system (the number of cointegrating vectors) is found by a procedure which is a kind of multivariate generalisation of the augmented Dickey-Fuller test. The following model is estimated:

$$
\begin{bmatrix}
\Delta r_t \\
\Delta e_t
\end{bmatrix} = 
\begin{bmatrix}
\mu_1 \\
\mu_2
\end{bmatrix} + 
\begin{bmatrix}
\gamma_{11} & \gamma_{12} \\
\gamma_{21} & \gamma_{22}
\end{bmatrix} 
\begin{bmatrix}
\Delta r_{t-1} \\
\Delta e_{t-1}
\end{bmatrix} + 
\sum_{i=1}^{3} 
\begin{bmatrix}
\Gamma_{1i} \\
\Gamma_{2i}
\end{bmatrix} \Delta f_{t-i}
- 
\begin{bmatrix}
\Pi_{11} & \Pi_{12} \\
\Pi_{21} & \Pi_{22}
\end{bmatrix} 
\begin{bmatrix}
\epsilon_{t-2} \\
\epsilon_{t-2}
\end{bmatrix} + 
\begin{bmatrix}
\epsilon_{1t} \\
\epsilon_{2t}
\end{bmatrix}
$$

where $[\epsilon_{1t}, \epsilon_{2t}]$ is a vector of Gaussian error term. The presence of lagged values of $\Delta f$ in the model is useful in coping with autocorrelation in the residuals without enlarging to much the lag length of the system. The number of stationary relationships (cointegration vectors) is given by the rank $r$ of the $\Pi$ matrix of the long-run coefficients. Once the number of stationary relationships is determined and their component estimated, standard asymptotic theory can be used for testing on the cointegrating vectors and analysing the shape of the error correction terms (e.g., testing to what extend receipts and expenses tend to be balanced in the long-run).

In our case, if $r = 1$, the long-run relationship is unique and the $\Pi$ matrix can be rewritten as $\alpha \beta'$ where $\beta$ is the $2 \times 1$ cointegrating vector and $\alpha$ the vector of the error correction coefficients. They measure the extent to which both variable adjust to the long-run.

Table 3 presents the cointegration tests used to determine $r$, the number of cointegration vectors. Two statistics are provided: the so-called maximal eigenvalue test and the trace test (see Johansen and Juselius (1990)). The two last columns give the threshold above which the null hypothesis is rejected at 95% and 90% respectively.

<table>
<thead>
<tr>
<th>(eigenvalue)</th>
<th>null</th>
<th>alternative</th>
<th>statistics</th>
<th>95% threshold</th>
<th>90% threshold</th>
</tr>
</thead>
<tbody>
<tr>
<td>$r = 0$</td>
<td>$r = 1$</td>
<td>14.74</td>
<td>14.90</td>
<td>12.91</td>
<td></td>
</tr>
<tr>
<td>$r \leq 1$</td>
<td>$r = 2$</td>
<td>6.77</td>
<td>8.18</td>
<td>6.50</td>
<td></td>
</tr>
</tbody>
</table>

<table>
<thead>
<tr>
<th>(trace)</th>
<th>$r = 0$</th>
<th>$r \geq 1$</th>
<th>21.51</th>
<th>17.95</th>
<th>15.66</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>$r \leq 1$</td>
<td>$r = 2$</td>
<td>6.77</td>
<td>8.18</td>
<td>6.50</td>
</tr>
</tbody>
</table>

The trace test allows us to conclude that there is only one cointegration relationship in the model both at 95% and 90% while the eigenvalue test gives the same conclusion only at the 90% threshold. This illustrates the fact that the two test procedures do not necessarily give the same result. The estimated eigenvalues used to compute the two test are respectively 0.126 and 0.060. It should be noted that these results reinforce our discussion of the previous section where we assumed our series to be I(1).4

4The restrictions that the variables alone are stationary are all rejected by the data. Moreover, if $\epsilon$ and $r$ were both I(0), then we would have found two cointegrating vectors.
The statistics on the residuals are presented in Table 4. Clearly, the absence of autocorrelation of order 1 to 5 cannot be rejected using Box-Pierce test. On the other hand, the Bera-Jarque test rejects the hypothesis of normality of the residuals. Whether the non-normality has important implications in terms of the determinations of the shape and the size of the cointegrating space is not clear-cut. As pointed out for example by Eitrheim (1992), non-normality does not affect the size and power of cointegration tests significantly. On the other hand, Franses and Haldrup (1992) pointed out that the presence of large additive outliers can give the impression that a given series is more stationary. As a consequence non-normality due to additive outliers can sometimes produce spurious I(0) series or spurious cointegrating vectors. While this might explain the outcomes of the unit root tests, we do not think it has an important effect on the outcome of the cointegration test (see Figure 4).

| Table 4 |

Residuals

<table>
<thead>
<tr>
<th>equation</th>
<th>BP(5)</th>
<th>95% threshold</th>
<th>Normality</th>
<th>95% threshold</th>
</tr>
</thead>
<tbody>
<tr>
<td>r</td>
<td>2.50</td>
<td>11.07</td>
<td>395.28</td>
<td>5.99</td>
</tr>
<tr>
<td>e</td>
<td>5.33</td>
<td>11.07</td>
<td>359.67</td>
<td>5.99</td>
</tr>
</tbody>
</table>

We try to reduce the model by imposing that the constant term $\mu$ appears only as part of the cointegrating vectors, so that there is no deterministic trend term in the variables. This model writes:

$$
\begin{bmatrix}
\Delta r_t \\
\Delta e_t
\end{bmatrix} =
\begin{bmatrix}
\gamma_{11} & \gamma_{12} \\
\gamma_{21} & \gamma_{22}
\end{bmatrix}
\begin{bmatrix}
\Delta r_{t-1} \\
\Delta e_{t-1}
\end{bmatrix}
+ \sum_{i=1}^{3}
\begin{bmatrix}
\Gamma_{1i} \\
\Gamma_{2i}
\end{bmatrix} \Delta f_{t-i}

- \begin{bmatrix}
\Pi_{11} & \Pi_{12} & \Pi_{13} \\
\Pi_{21} & \Pi_{22} & \Pi_{23} \\
\Pi_{31} & \Pi_{32} & \Pi_{33}
\end{bmatrix}
\begin{bmatrix}
r_{t-2} \\
e_{t-2} \\
1
\end{bmatrix}
+ \begin{bmatrix}
\epsilon_{1t} \\
\epsilon_{2t}
\end{bmatrix}
$$

(4)

Table 5 displays the cointegration tests for this model. The corresponding eigenvalues are: 0.129 and 0.061. Leading to the same conclusion as above.

| Table 5 |

Dimension of the cointegration space (case with a restricted constant)

<table>
<thead>
<tr>
<th>(eigenvalue)</th>
<th>null</th>
<th>alternative</th>
<th>statistics</th>
<th>95% threshold</th>
<th>90% threshold</th>
</tr>
</thead>
<tbody>
<tr>
<td>$r = 0$</td>
<td>$r = 1$</td>
<td>15.10</td>
<td>15.67</td>
<td>13.75</td>
<td></td>
</tr>
<tr>
<td>$r \leq 1$</td>
<td>$r = 2$</td>
<td>6.85</td>
<td>9.24</td>
<td>7.52</td>
<td></td>
</tr>
<tr>
<td>(trace)</td>
<td>$r = 0$</td>
<td>$r \geq 1$</td>
<td>21.95</td>
<td>19.96</td>
<td>17.85</td>
</tr>
<tr>
<td>$r \leq 1$</td>
<td>$r = 2$</td>
<td>6.85</td>
<td>9.24</td>
<td>7.52</td>
<td></td>
</tr>
</tbody>
</table>

The unique cointegration vector, normalized to have a unit coefficient for $r$, implies the following long-run relationship:

$$
r = 0.70 + 1.25 e
$$

(5)

This can be compared to what we obtain using the Engle and Granger (1987) method which amounts to estimate the long-run relationship by static OLS regression:

$$
r = 1.50 + 0.95 e
$$

(6)
with

\[ CRDW = 48, \ DF = -3.56 \ ADF(1) = -3.71, \ ADF(2) = -4.88 \]

where \( CRDW \), \( DF \) and \( ADF(.) \) are the conventional residual-based cointegration tests proposed by Engle and Granger (1987). No cointegration is rejected by all these tests at 5%.

In order to get a better understanding of the long-run relationship, we first test for the exclusion from the long-run relationship of each variable individually which basically amounts to test for the stationarity of the remaining variable. These two assumptions are rejected. The last hypothesis we test is a unit long-run elasticity between expenditures and receipts which amounts to test the stationarity of the ratio of the receipts to the expenditures. This last hypothesis is not rejected at the 5% level.

<table>
<thead>
<tr>
<th>Hypothesis</th>
<th>Interpretation</th>
<th>test</th>
<th>95% threshold</th>
</tr>
</thead>
<tbody>
<tr>
<td>((\beta_0, 0, \beta_2))</td>
<td>( r ) is ( I(0) )</td>
<td>6.75</td>
<td>3.84</td>
</tr>
<tr>
<td>((\beta_0, \beta_1, 0))</td>
<td>( e ) is ( I(0) )</td>
<td>8.21</td>
<td>3.84</td>
</tr>
<tr>
<td>((\beta_0, -1, 1))</td>
<td>( r - e ) is ( I(0) )</td>
<td>1.41</td>
<td>3.84</td>
</tr>
</tbody>
</table>

The resulting long-run relation has the form:

\[ r = 1.10 + 1 \ e \]  

(7)

The corresponding deviation from the long-run equilibrium (7) is graphed in Figure 4.

Figure 4: Deviation from the long run equilibrium (7)

Johansen’s procedure also allows to test the presence of cointegrating vectors in each equation of the system, using a likelihood ratio test whose critical value is distributed as
An interesting thing about this result is its interpretation in terms of causal ordering and weak exogeneity (see Engle et al. (1983)). First, we see that when the cointegrating vector does not enter the equation of the receipts, these do not react to lagged levels of expenditures so that one can conclude that expenditures do not Granger cause receipts. However, when the cointegrating vector is sufficiently exogenous in terms of the potential efficiency of a partial model, the implication in the equation for the receipts is the implication in terms of exogeneity, i.e., the implication in the cointegrating vector is sufficient to ensure that receipts can be treated as weakly exogenous for the parameters of the long-run relation. This enables us to eventually focus on the conditional model. This will allow us to derive a more meaningful short-run dynamic model whose parameters will easily be determined and which will be more easily reparametrized and reduced.

### Table 7

<table>
<thead>
<tr>
<th>Absence of cointegrating vectors</th>
<th>Equation Test</th>
<th>95% Threshold</th>
<th>3.84</th>
</tr>
</thead>
<tbody>
<tr>
<td>r</td>
<td>0.25</td>
<td>6.61</td>
<td></td>
</tr>
<tr>
<td>c</td>
<td></td>
<td>3.84</td>
<td></td>
</tr>
</tbody>
</table>

The analysis of a medieval judicial system confirms the hypothesis of causality from expenditures to the activities of the serjeants, in proportion of the fines. The receipts result from the collection of fines. In all cases, these expenditures are directly linked to the activity of the serjeants, in proportion of the fines. The receipt collection activity is characteristic of a "tax-oriented" system which seems therefore fully compatible with our data.

**Testing cointegration in a conditional model**

As just pointed out, from the previous analysis, it appears that our variables are cointegrated with a unique cointegrating vector entering only the equation for the expenditures. In this subsection, we now try to analyse the same data in a conditional model where the dynamic of the expenditures are retained as the only endogenous variable. The numbers of fines (f) and the receipts (r) are maintained as exogenous in a first stage of the analysis. The particular advantages of such a partial (conditional) approach are, at least, under the null hypothesis of weak exogeneity, the approach is equivalent to maximum likelihood estimation and hence inference on the long run parameters can be conducted by means of traditional $\chi^2$ statistics; it allows economically meaningful restrictions to be

---

5. Brunner and Nicolai (1992) talk about persistent (non)causality in opposition to (short run) Granger (non)causality.
easily imposed on the long-run parameters while further overidentifying restrictions can also be introduced at a later stage.

We consider a simple single equation error correction model (ECM):

$$\Delta e_t = \kappa_0 + \sum_{i=0}^{p} \kappa_1 \Delta r_{t-i} + \sum_{i=0}^{s} \kappa_2 \Delta f_{t-i} + \sum_{i=1}^{q} \kappa_3 \Delta e_{t-i} + \kappa_4 T$$

$$-\lambda \begin{bmatrix} \beta_1 & \beta_2 & \beta_3 \end{bmatrix} \begin{bmatrix} r_{t-1} \\ e_{t-1} \\ f_{t-1} \end{bmatrix} + \eta_t$$

(8)

where $f$ and $r$ are weakly exogenous variables and $\eta_t$ is a Gaussian i.i.d. error-term. In this single equation context, and given the Granger Representation Theorem (Engle and Granger (1987)), cointegration will hold as far as the error correction term is significantly different from zero. Boswijk (1991a) therefore proposes to test for the significance of the cointegrating relationships using a simple Wald test for the significance of the error correction term, based on OLS estimation of the model. This way of testing for the presence of cointegrating relationships is a natural generalisation of the simple $t-test$ on the ECM term proposed by Kremers et al. (1992) for the case of bivariate relationships with known cointegration relations. The limiting distribution of the Wald test is not of $\chi^2$ type and is expressed as a functional of a vector Brownian motion whose critical values are tabulated in Boswijk (1991a). The estimates of the long run coefficients are easily obtained by using indirect estimators. The properties of the Wald test and the Indirect Least Squares estimation (for a unique cointegrating vector) seem reliable in the simulation outcome reported by Boswijk and Franses (1992), at least compared to the Engle and Granger (1987) two-step approach and Johansen's maximum likelihood approach. Although the approach is initially developed for unrestricted conditional ECMs, we may clearly expect finite sample gains by parsimoniously reparameterising the short run dynamic of the model.

The application of this approach leads to the following ECM where a trend term is added to allow for the observed trend stationary character of $f$ as well to ensure the similarity of the Wald statistics (see Boswijk (1991a)):

$$\Delta e_t = -2.3 + .007 \quad T - 1.03 \quad D1451 + .86 \quad \Delta r_t + .33 \quad \Delta r_{t-3}$$

$$(.5) \quad (.001) \quad (.23) \quad (.056) \quad (.082)$$

$$- .33 \quad \Delta e_{t-3} - .19 \quad \Delta e_{t-6} - .41 \quad \Delta f_{t-1} - .36 \quad \Delta f_{t-2} - .28 \quad \Delta f_{t-3}$$

$$(.074) \quad (.045) \quad (.095) \quad (.089) \quad (.103)$$

$$+ .17 \quad r_{t-1} - .24 \quad e_{t-1} + .59 \quad f_{t-1}$$

$$(.054) \quad (.062) \quad (.113)$$

(9)

with

$$R^2 = .80 \quad DW = 2.07 \quad WALD = 31.44 \quad SER. COR. = 0.23 \quad NORM = 0.64 \quad HET = 1.75$$

SER.COR. is the Lagrange multiplier test of residual serial correlation of order one (distributed as a $\chi^2_1$), NORM is the Bera-Jarque normality test$^6$ (distributed as a $\chi^2_2$).

---

$^6$The inclusion of the dummy D1451 is necessary to have normality of the residuals.
and HET is the heteroscedasticity test based on the regression of squared residuals on squared fitted values (distributed as a $\chi^2$). WALD is a Wald test for the significance of the error correction term, under the null of no-cointegration its distribution is tabulated in Boswijk (1991a). The 5% critical value equals 19.17 so that we can safely reject the null hypothesis of no-cointegration. The estimated long run relation obtained using the ILS estimator reads as

$$r = 13.6 + 1.43 e - [3.46 f + 0.044 T]$$

(10)

The corresponding deviation from the long run equilibrium is graphed in Figure 5 and can be compared to the Johansen's one in Figure 4.

Figure 5: Deviation from the long run equilibrium (10)

As pointed out in Boswijk (1991b) and Urbain (1992) it is also possible to test for the weak exogeneity character of the receipts in this framework by fitting an unrestricted (marginal reduced form) ECM for $\Delta r_t$ and then testing the significance of the error correction term which was estimated at the previous step. The Wald test for this null hypothesis (in $\chi^2$ form) gives a computed value of 2.37, well below the 5% critical value which equals 3.84. This results clearly confirms the results obtained within the Johansen framework. Consequently, efficient inference on the long-run relationship does not require knowledge of the way receipts are generated.

\footnote{However, as pointed out in Urbain (1993) this condition may not be sufficient for weak exogeneity if the parameters of interest also include the short run dynamic of the model, especially when this has been reparametrised. In that case a usual orthogonality condition between the potential exogenous and the error term has also to be jointly satisfied. This last hypothesis can be easily tested by extending usual orthogonality tests (see for example Pesaran and Smith (1990)). The framework of expanded regression is particularly appealing in this case. All what needs to be done is to estimate a (single equation structural) ECM for $\Delta e_t$ and test jointly for the significance of the estimated residual of this model and of the cointegrating vector in a marginal reduced form ECM for $\Delta r_t$ using usual variable addition test. The Wald statistic for this hypothesis, distributed as a $\chi^2(2)$ under the null hypothesis of weak exogeneity (w.r.t. both the short and long run parameters) gives a a value of 2.737 well below the 5% critical value.}
We can also use this marginal ECM for $\Delta r_t$ to test the null hypothesis of Granger non-causality from $e_t$ to $r_t$, by testing jointly the absence of all lagged $\Delta e_t$ and the absence of the cointegrating vector (see Phillips and Toda (1992)). The Wald test for this Granger non-causality null hypothesis gives a computed value of 5.04 ($\chi^2(5)$) so that the hypothesis of Granger non-causality from expenditures to receipts cannot be rejected by the data. As a consequence, we are now able to significantly reduce the size of the problem, a fact which will be quite useful in the next section since receipts may now be treated as fixed for the purpose of inference on the long-run relationship.

Another implication of these exogeneity and causality results is that we can use standard asymptotic distribution to conduct inference on the long-run relationship. Testing for example the restriction implied by a long run unit elasticity of receipts to expenditures is rejected at the usual 5% level (computed statistics of 4.60 $\chi^2(1)$), contradicting the result obtained with Johansen’s procedure.

While this is a direct statistical implication of the previous results, this exogeneity and causality outcome has also some interesting behavioral content (see the discussion in Richard (1980)). In particular, the non rejection of weak exogeneity of receipts for the parameters of the expenditure model implies that the relation between expenditures, receipts and fines as modelled above will not be modified by an unexpected change in the receipts. As we know that the receipts have nevertheless undergone important changes this could eventually explain the breakdown of the system. However, this will be implicitly analysed in the next subsection.

Figure 6: Dynamic simulation of the marginal model

4 The Stability of Receipt Collection

If we are already able to point out some interesting facts about the relation between receipt and expenditures in Nivelles’ town over our sample period, our analysis up to
now does not enable us to really assess whether our model has really isolated an invariant relation between our variables. In particular, if we want to be ensured that the model we have estimated and formulated for the expenditures remains valid even under various receipts regimes, we should be ensured that receipts are super exogenous (in the sense of Engle et al., 1983) so that fundamental changes in the receipts setting process would not alter the way expenditures are linked to the remaining variables. Although explicit tests for super exogeneity have been recently proposed by Engle and Hendry (1993), we will rely on an alternative way for assessing the empirical validity of our model. We must first again complete the short term model (9) with an equation for the receipts (the marginal model). The latter is formulated as a function of the number of fines. Given the outcome of weak exogeneity and non-causality tests obtained in the previous section, a general to specific approach leads to the following specification:

$$\Delta r_t = 0.84 \Delta f_t - 0.16 \Delta f_{t-2} + 0.06 \Delta r_{t-1} + 0.15 \Delta r_{t-2} - 0.08 \Delta r_{t-3}$$  

$$+ 2.24 D_{1437}$$

(0.066) (0.084) (0.048) (0.06) (0.047)  

(0.22)

$$R^2 = .77 \quad DW = 1.85 \quad SER.COR. = 0.66 \quad NORM = 18.22 \quad HET = 0.54$$

The dummy variable in 1437 takes into account the normative text from the Duke of Brabant which has increased the amount of the fines in Nivelles by 100%. Unfortunately, the inclusion of this variable is not sufficient to warrant normality of the residuals.

Figure 7: Dynamic simulation of the conditional model

The dynamic behaviour of the system is then investigated using dynamic simulations of (9) and (11) independently (Figures 6 and 7) and then jointly (Figure 8). The underlying motivation is simple: if receipts are effectively super exogenous for the parameters of (9)
and not Granger caused by expenditures, then the conditional model (9) should be able to simulate well the true evolution of the expenditures. The independent simulations show that the conditional model performs quite well while the marginal model displays a positive bias at the end of the period. If we now simulate jointly the two models, this bias deteriorates the performance of the model for $e$ in the joint simulation. This points out that although the way receipts are generated has undergone at least one important change, the conditional model alone tracks surprisingly well the observed data.

This is confirmed by a recursive within sample analysis which points out the overall within sample stability of the conditional model although the marginal model for the receipts displays instability. In particular, we observe that the marginal model (receipts) displays instability around 1520: the elasticity of receipts of fines increases (since fines are declining, receipts are declining faster after 1520).

Figure 8: Dynamic simulation of the full model

![Graph showing the dynamic simulation of the full model.](image)

The interest of having derived a conditional ECM with valid conditioning variables is that the instability in receipt collection is limited to its source, i.e. it does not affect the relation between receipts and expenditures. We have been able to derive a model for the expenditures which isolates the invariant and constant features of the phenomena we analyse. The implications for the breakdown of the system are now clearer: Although the relation between expenditures and receipts has remained the same over the sample period, the link between receipts and fines has undergone a substantial change in 1520. This change can be related to the four following elements.

A first element could a change in the structure of fines. In general, the fines are composed of fixed fines (called "faits mandés") and by pilgrimage fines. These pilgrimage fines are initially composed of bail payments to avoid the sentence of having to go to a designated place of pilgrimage. The distribution between the two types of fines is not the same in the 15th and in the 16th century: we observe a decreasing proportion of fixed value fines and a growing proportion of pilgrimage fines in the 16th century. An explanation can therefore be found in the decrease of "good" receipts (fixed fines) and
the increase of less profitable fines (pilgrimages). A pilgrimage is less profitable because it can be used as physical punishment by order of the authorities (see Van Herwaarden (1978)) or by the offender’s choice, so that an actual journey will often replace a receipt flow.

Table 8
Structure of Fines

<table>
<thead>
<tr>
<th></th>
<th>1457-1500</th>
<th>1500-1537</th>
</tr>
</thead>
<tbody>
<tr>
<td>Fixed fines</td>
<td>64%</td>
<td>52%</td>
</tr>
<tr>
<td>Pilgrimages</td>
<td>15%</td>
<td>28%</td>
</tr>
<tr>
<td>Others</td>
<td>21%</td>
<td>20%</td>
</tr>
</tbody>
</table>

Second, within the category of the fixed fines, the structure also changed between the 15th and the 16th century. The highest fines (10, 20 and 60 “oboles”) declined from 47% to 30% of fixed fines, while the lower ones (1 to 5 “oboles”) have increased from 53% to 70%.

Table 9
Structure of fixed fines

<table>
<thead>
<tr>
<th></th>
<th>1457-1500</th>
<th>1500-1537</th>
</tr>
</thead>
<tbody>
<tr>
<td>60 “oboles”</td>
<td>2 %</td>
<td>0%</td>
</tr>
<tr>
<td>20 “oboles”</td>
<td>44 %</td>
<td>24%</td>
</tr>
<tr>
<td>10 “oboles”</td>
<td>1 %</td>
<td>6%</td>
</tr>
<tr>
<td>4 “oboles”</td>
<td>49 %</td>
<td>64%</td>
</tr>
<tr>
<td>2 “oboles”</td>
<td>3 %</td>
<td>4%</td>
</tr>
<tr>
<td>1 “oboles”</td>
<td>1 %</td>
<td>2%</td>
</tr>
</tbody>
</table>

Third, Rousseaux (1990a) shows that the proportion of unperceived fines as well as the proportion of unsentenced offenders tends to grow.
Fourth, a detailed analysis of the accounts for the years 1517-1536 reveals a change in the vocabulary: new terms related to corporal punishments are introduced. (Lord's justice becomes Criminal Justice). This change in vocabulary points to a change in the nature of penalties, implying that the proportion of monetary sanctions went down during the 16th century, as illustrated in Table 10. Table 10 is also consistent with the fact that the emerging modern judicial system requires more formal proof of the offenses, leading to a more intensive use of torture and physical punishment (see Langbein (1974) and (1977)).

<table>
<thead>
<tr>
<th></th>
<th>15c. Number</th>
<th>%</th>
<th>16c. Number</th>
<th>%</th>
</tr>
</thead>
<tbody>
<tr>
<td>Monetary sanctions</td>
<td>4636</td>
<td>97.6</td>
<td>979</td>
<td>76.4</td>
</tr>
<tr>
<td>Corporal sanctions</td>
<td>113</td>
<td>2.4</td>
<td>302</td>
<td>23.6</td>
</tr>
</tbody>
</table>

To summarize, the instability in the link between receipts and fines can be explained by the evolution of the structure of receipts and of the number of fines. The following elements are worth noting:

1. The mean value of a fine decreased from the 15th to the 16th century.
2. To some extent, this evolution is amplified by the increase of unperceived fines.
3. Monetary sanctions declined, while physical sanctions grew.
4. The judicial pilgrimage played an important role in the transformation from financial into physical sanctions.

**Conclusion**

In this paper we have used recently developed time series techniques to analyse long historical data series on the budget of the Court of Nivelles (Belgium) in the late Middle Ages.

Our major empirical findings point out (1) a deterministic negative trend in the number of fines; (2) stochastic trends in receipts and expenditures; (3) the existence of a balanced long-term relationship between expenses and receipts; (4) unidirectional causality from receipts to expenditures: This supports the idea that the medieval system under consideration is a fine-focused (tax-oriented) control system; (5) dynamic simulations and exogeneity analysis point out that although the way receipts are generated has undergone at least one important change, the relationship between receipts and expenditures has remained constant over time. In particular, we observe that the elasticity of receipts to fines increases around 1520 (since fines are declining, receipts are declining faster). This change in the "productivity" of fines can probably be related to some historical elements such as the decline in the number of monetary sanctions along with the growth of physical sanctions (The judicial pilgrimage played an important role in this transformation from financial into punitive sanctions). Our empirical evidence on the presence of changes in receipt collection around 1520 favours the assumption of a transformation of a medieval, urban-based, tax-oriented system into a modern, State-based, punishment-focused system.
References


