

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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Time-series techniques are used 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 this mediaeval system was 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 mediaeval, urban-based, tax-oriented system and its transformation into a system based on punishment.

I. INTRODUCTION

The Middle Ages are frequently considered as a period of obscurantism and used to be well known as the ‘dark ages’. In fact, mediaeval 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. These reveal a fully 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, Gheldres, etc. The accounts of local, regional and principal officers provide us with substantial information about the financing of the administration of justice.

Rousseaux (1990) analyses various data series on the costs and benefits of justice in the late Middle Ages and the early modern period, 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

stabilization and effective functioning of this taxation system from 1457–1520; third, a crisis in the system and its progressive collapse between 1520–1536. From these observations, based on the accounts of the officer of Nivelles, Rousseaux (1990) 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 here is to provide some empirical elements on (1) the existence of a mediaeval ‘tax-oriented’ system in Nivelles and (2) its progressive change towards a more ‘punishment-oriented’ system around 1520–1530.

The first point 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. 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, this 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.

The second point is 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 history¹ have focused on this period (exceptions are Muchembled, 1992 and Chiffolleau, 1974). Those who have centred their interest on this period – for example, Brouwers (1965) and Marounek (1980) – do not succeed in proposing a precise chronology of the mutation.

This mutation took place in a deeply changing context in European societies at the end of the 15th and the beginning of the 16th centuries. From an *economic* point of view, the general situation deteriorated at the end of the 15th 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 between 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 14th–15th centuries, based on textile and cloth industries, was challenged around 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 Western societies. The Protestant reformation led to a crisis in social relationships, especially in the Netherlands. Lutherans, and later Anabaptists and Calvinists, were prosecuted by the Spanish Kings Charles V and Philip II from 1520 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 16th century was characterized 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 V signed a large number of edicts: the 'placards' against the heretics. This was a great innovation. For the first time, general criminal legislation was promulgated in all the Low countries (see Rousseaux, 1990 and 1995 and Muchembled, 1992). Part of this legislation appeared in the local records of the judges of Nivelles. At the same time, legists such as 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 play 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 II 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 population 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, p 133). 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. Here, their hypothesis can be related to the shift, between the 15th and the 16th centuries, from justice by taxation (fines) to punitive justice.

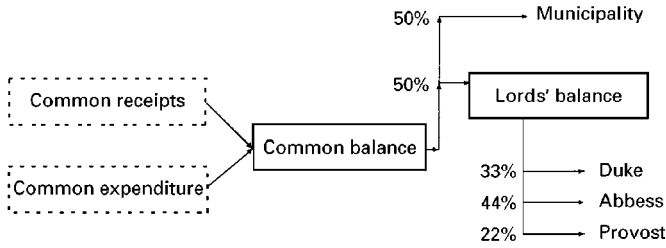
Besides these political, religious and economic explanations we want to test econometrically the potential occurrence of the structural break implied by the second point above. At this stage, the exact nature of the structural break is unknown: we do not know if the mutation of the system should affect 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 (super) 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.

Section II presents the data sources and the basic functioning of the budget of the Court of Nivelles. In Section III, the statistical properties of the data are analysed. The long-run relationship between receipts and expenditures and their causality links are investigated in Section IV. In Section V we address the problem of structural stability of the model and its link with the collection of fines. Section VI concludes.

II. 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

¹ General overviews are in Cockburn (1977), Brewer and Styles (1980), Soman (1980), Gattrell *et al.* (1980); see also Dupont-Bouchat (1976) for Belgium and Weisser (1982) for a broad synthesis.

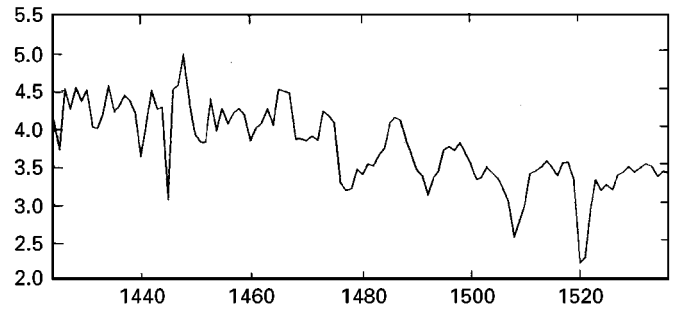
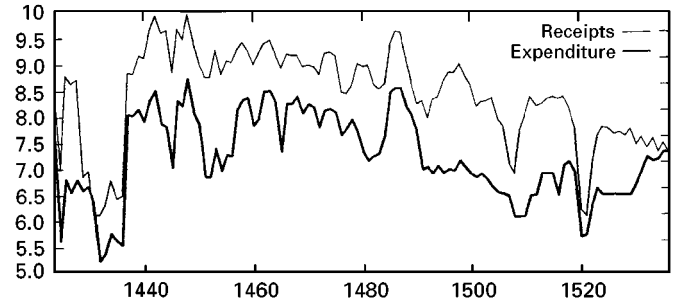
Fig. 1. *The court budget*

(14th–18th centuries). 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.

The way in which the various receipts and expenditures are shared among the authorities is presented in Fig. 1. Common receipts (fines) and common expenditures (preventive imprisonment and management of justice) are allocated locally 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 1423–1550. These accounts are complete for the period 1423–1536, i.e. for more than 100 years. There were three problems in homogenizing the series:

1. *The diversity of the structure of the accounts:* The internal structure of this accountancy changed over this long period. The structure is unstable from 1423–1457. 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 (Fig. 2) and common receipts and expenditures (Fig. 3).
2. *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. These series are homogenized using a three-period moving average.
3. *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’). As a consequence of the fluctuating value of monetary units, these relations changed over time. For the analysis, all data are expressed in plaques (1 plaque = 1.20 pounds = 0.33 florin). Between 1438–1536, this unit of account seems to remain stable.

Fig. 2. *Number of fines*Fig. 3. *Common budget*

III. STATISTICAL PROPERTIES OF THE DATA

As it is clear from Figs. 2, 3 and 4, the data series display trending behaviour. Consequently, before searching for stable relationships between the series, we have to test whether or not they are trend stationary or integrated processes. This is an 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 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 non-stationarity in the data, we first investigate the integration order of the series using two different unit root tests: 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:

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

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 $\rho = 1$, the shocks ε_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 Equation (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 ε_t . In practice, it may be more realistic to allow for some dependence among the ε_t 's (e.g. allowing for serial correlation and heteroskedasticity, but also controlling the extent 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 for even higher-order trend polynomials.

From the figures reported above, it appears clear that most of the series are likely to possess some deterministic trend component, or might even be characterized 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 1. For both ADF-type and Phillips-Perron type statistics, the lag truncation was set at four, after having scrutinized the autocorrelation function of the first-differenced processes. The following notation is used: τ is the usual t-test based on (Equation 1) without trend term, τ_μ is the t -test based on (Equation 1) with trend term. The Φ_i , $i = 1, 2, 3$ are one-sided F-type statistics respectively testing the null hypotheses of (a) $(\mu, \rho) = (0, 1)$ in Equation (1) without trend, (b) $(\mu, \alpha, \rho) = (0, 0, 1)$ in Equation (1) with trend, and (c) $(\mu, \alpha, \rho) = (\mu, 0, 1)$ in (Equation 1) with trend. The 'Z' versions reported in Table 1 have the same interpretation but are computed with the Phillips-Perron non-parametric correction instead of the lag augmentation.²

Table 1. Unit root tests – levels

Test statistic	r	e	f	95% Critical value
τ	-2.38	-2.73	-1.94	-2.89
τ_μ	-2.77	-3.11	-3.96	-3.45
Φ_1	2.87	3.74	2.04	4.71
Φ_2	2.67	3.26	5.35	4.88
Φ_3	3.98	4.88	7.87	6.49
$Z(\tau)$	-2.83	-3.36	-3.27	-2.89
$Z(\tau_\mu)$	-3.48	-4.70	-5.74	-3.45
$Z(\Phi_1)$	4.04	5.65	5.36	4.71
$Z(\Phi_2)$	3.34	4.05	11.02	4.88
$Z(\Phi_3)$	4.99	6.07	16.50	6.49

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

As shown in Table 2, we reject the null of a unit root with drift against the trend stationary alternative hypothesis for f . The same battery of tests applied to the two remaining series confirms, at least at this stage, that the 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 characterized as $I(0)$ processes. This phenomena could eventually be explained by various factors, including how the data were constructed. However, the point estimates of the first-order autoregressive coefficient are, for r and e near unity. We shall, therefore, follow the advice of among others, Campbell and Perron (1991)³ and consider r and e as integrated processes of the order of one.

Any attempt to build conditional or simultaneous equation models linking these different variables should, thus, clearly recognize the individual properties of the data. In particular, models linking common receipts and expenditures should be built by means of cointegration techniques explicitly to take the stochastic nature of their trending 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 periods 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 for 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 increased the amount of fines in Nivelles by 100%. Before 1438, one fine of one obole (real unit money)

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 four.

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.

Table 2. Dimension of the cointegration space (case with a general constant)

	Null	Alternative	Statistics	95% critical value	90% critical value
(eigenvalue)	$r = 0$	$r = 1$	14.74	14.90	12.91
	$r \leq 1$	$r = 2$	6.77	8.18	6.50
(trace)	$r = 0$	$r \geq 1$	21.51	17.95	15.66
	$r \leq 1$	$r = 2$	6.77	8.18	6.50

was evaluated to 21 plaques 2 gros. After this date, one obole was valued as 42 pl. 4 gros. The break is also due to a structural reorganization of judicial taxation. In the years up to 1520, receipts tend to decline faster in contrast with the appearance of physical punishment, registered not in receipt but in expenditure, given the costs of physical executions (for example, prison, torture and hanging). However, the theoretical developments of segmented trend functions with multiple breaks is yet not well developed, so this is left for further research.

IV. LONG-RUN RELATIONSHIPS

This section focuses on receipts and expenditures of the common budget, and we investigate if, although these variables are non-stationary, there exists 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 to the 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 a linear combination, this can be interpreted as a long-run relationship that holds between the variables over more than one century, in the sense that 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.

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 generalization of the augmented Dickey–Fuller test. The

following model is estimated:

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

where $[\varepsilon_{1t} \varepsilon_{2t}]'$ is a vector of Gaussian error term. The presence of lagged values of Δf in the model is useful to cope with autocorrelation in the residuals without enlarging too much the lag length of the system. The number of stationary relationships (cointegration vectors) is given by the rank r of the Π 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 extent receipts and expenses tend to be balanced in the long run).

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

Table 2 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 last two columns give the threshold above which the null hypothesis is rejected at 95% and 90% respectively. 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 tests are respectively 0.126 and 0.060. It should be noted that these results reinforce the discussion above where the series was assumed to be $I(1)$.⁴

⁴ The restrictions that the variables alone are stationary are all rejected by the data. Moreover, if e and r were both $I(0)$, then we would have found two cointegrating vectors.

The statistics on the residuals are presented in Table 3. Clearly, the absence of autocorrelation of order 1–5 cannot be rejected using a 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 clearcut. 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 Fig. 4).

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

$$\begin{aligned} \begin{bmatrix} \Delta r_t \\ \Delta e_t \end{bmatrix} &= \begin{bmatrix} \gamma_{11} & \gamma_{12} \\ \gamma_{21} & \gamma_{22} \end{bmatrix} \begin{bmatrix} \Delta r_{t-1} \\ \Delta e_{t-1} \end{bmatrix} + \sum_{i=1}^3 \begin{bmatrix} \Gamma_{1i} \\ \Gamma_{2i} \end{bmatrix} \Delta f_{t-i} \\ &- \begin{bmatrix} \Pi_{11} & \Pi_{12} & \Pi_{13} \\ \Pi_{21} & \Pi_{22} & \Pi_{23} \\ \Pi_{31} & \Pi_{32} & \Pi_{33} \end{bmatrix} \begin{bmatrix} r_{t-2} \\ e_{t-2} \\ 1 \end{bmatrix} + \begin{bmatrix} \varepsilon_{1t} \\ \varepsilon_{2t} \end{bmatrix} \end{aligned} \quad (4)$$

Table 4 displays the cointegration tests for this model. The corresponding eigenvalues are: 0.129 and 0.061, leading to the same conclusion as above.

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

$$r = 0.70 + 1.25 e \quad (5)$$

This can be compared to what is obtained using the Engle and Granger (1987) method, which amounts to estimating the long-run relationship by static OLS regression:

$$r = 1.50 + 0.95 e \quad (6)$$

with

$$CRDW = 0.48, \quad DF = -3.56$$

$$ADF(1) = -3.71, \quad ADF(2) = -4.88$$

where $CRDW$, DF and $ADF(\cdot)$ 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 arrive at a better understanding of the long-run relationship, we first test for the exclusion from the long-run relationship of each variable individually. Basically, this amounts to testing for the stationarity of the

Table 3. *Residuals*

Equation	BP(5)	95% critical value	Normality	95% critical value
r	2.50	11.07	395.28	5.99
e	5.33	11.07	359.67	5.99

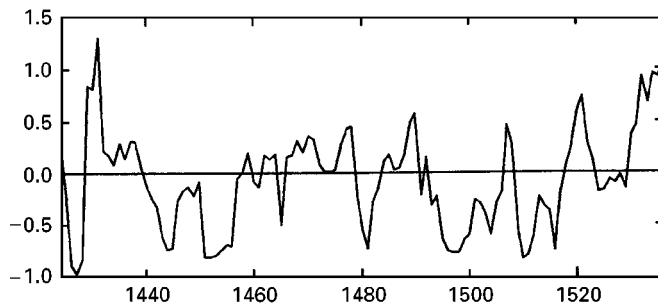


Fig. 4. *Deviation from the long-run equilibrium (Equation 7)*

remaining variable. These two assumptions are rejected. The last hypothesis tested is a unit long-run elasticity between expenditures and receipts which amounts to testing the stationarity of the ratio of the receipts to the expenditures. This last hypothesis is not rejected at the 5% level.

The resulting long-run relation has the form:

$$r = 1.10 + 1 e \quad (7)$$

The corresponding deviation from the long-run equilibrium (Equation 7) is shown in Fig. 4.

Johansen’s procedure also allows one to test for the presence of cointegrating vectors in each equation of the system, using a likelihood ratio test whose critical value is distributed as a χ^2 . This aims at analysing whether the variations of the variables adjust to the long run of the model. This test is presented in Table 6. Only the expenditures adjust to deviations from the long-run relationship.

One interesting factor about this result is its interpretation in terms of causal ordering and weak exogeneity (see Engle *et al.*, 1983). First, 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 ‘long-run’ Granger-cause receipts.⁵ Whether some short-run causality can nevertheless remain will be analysed later on. The second interesting consequence of the absence of the cointegrating vector in the equation for the receipts is the implication in terms of exogeneity – in other words, the implication in terms of the potential efficiency of a partial model. As pointed out *inter alia* by Johansen (1992) and Urbain (1992), in the framework of error-correction models, the absence of the cointegrating vector is sufficient to ensure that receipts can be treated as weakly exogenous for the parameters of the long-run

⁵ Bruneau and Nicolai (1992) talk about persistent (non) causality in opposition to (short-run) Granger (non) causality.

Table 4. Dimension of the cointegration space (case with a restricted constant)

	Null	Alternative	Statistics	95% critical value	90% critical value
(eigenvalue)	$r = 0$	$r = 1$	15.10	15.67	13.75
	$r \leq 1$	$r = 2$	6.85	9.24	7.52
(trace)	$r = 0$	$r \geq 1$	21.95	19.96	17.85
	$r \leq 1$	$r = 2$	6.85	9.24	7.52

Table 5. Shape of the cointegration vector

Hypothesis	Interpretation	Test	95% critical value
$(\beta_0, 0, \beta_2)$	r is $I(0)$	6.75	3.84
$(\beta_0, \beta_1, 0)$	e is $I(0)$	8.21	3.84
$(\beta_0, -1, 1)$	$r - e$ is $I(0)$	1.41	3.84

Table 6. Absence of cointegrating vectors

Equation	Test	95% critical value
r	0.026	3.84
e	6.61	3.84

relation. This enables us eventually to focus on a conditional model. This will allow us to derive a more meaningful short-run dynamic model whose parameters will be easier to interpret and which will be more easily reparameterized and reduced.

The analysis of the mediaeval judicial system confirms the hypothesis of causality from receipts to expenditures. The receipts result from the collection of fines. The expenditures are generally linked with the activity of collection (benefits for the mayor and sergeant, in proportion to the fines) and sometimes with the prison wardens' expenditures. In all cases, these expenditures are directly provoked by the receipt-collection activity. This unidirectional causality is a characteristic of a 'tax-oriented' system which seems, therefore, fully compatible with the data.

Testing cointegration in a conditional model

As discussed above, it appears that the variables are cointegrated, with a unique cointegrating vector entering only the equation for the expenditures. Here, we try to analyse the same data in a conditional model, where 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 threefold: under the testable 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 χ^2 statistics; it allows economically meaningful restrictions to be 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):

$$\Delta e_t = \kappa_0 + \sum_{i=0}^p \kappa_{1i} \Delta r_{t-i} + \sum_{i=0}^s \kappa_{2i} \Delta f_{t-i} + \sum_{i=1}^q \kappa_{3i} \Delta e_{t-i} + \kappa_4 T - \lambda [\beta_1 \ \beta_2 \ \beta_3] \begin{bmatrix} r_{t-1} \\ e_{t-1} \\ f_{t-1} \end{bmatrix} + \eta_t \tag{8}$$

where f and r are weakly exogenous variables and η_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 generalization 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 χ^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 reparameterizing 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):

$$\begin{aligned} \Delta e_t = & -2.3 + 0.007 T - 1.03 D1451 + 0.86 \Delta r_t \\ & (0.5) \quad (0.001) \quad (0.23) \quad (0.056) \\ & + 0.33 \Delta r_{t-3} - 0.33 \Delta e_{t-3} - 0.19 \Delta e_{t-6} \\ & (0.082) \quad (0.074) \quad (0.045) \\ & - 0.41 \Delta f_{t-1} - 0.36 \Delta f_{t-2} - 0.28 \Delta f_{t-3} \\ & (0.095) \quad (0.089) \quad (0.103) \\ & + 0.17 r_{t-1} - 0.24 e_{t-1} + 0.59 f_{t-1} \end{aligned} \quad (9)$$

with

$$\begin{aligned} R^2 = 0.80 \quad DW = 2.07 \quad WALD = 31.44 \\ SER.COR. = 0.23 \quad NORM = 0.64 \quad HET = 1.75 \end{aligned}$$

SER.COR. is the Lagrange multiplier test of residual serial correlation of order one (distributed as a χ_1^2), NORM is the Bera-Jarque normality test⁶ (distributed as a χ_2^2) and HET is the heteroscedasticity test based on the regression of squared residuals on squared fitted values (distributed as a χ_1^2). 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

$$r = 13.6 + 1.43 e - [3.46 f + 0.044 T] \quad (10)$$

The corresponding deviation from the long-run equilibrium is shown in Fig. 5 and can be compared to that of Johansen in Fig. 4.

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 Δr_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 χ^2 form) gives a computed value of 2.37, well below the 5% critical value which equals 3.84. This result clearly confirms the results obtained within the Johansen

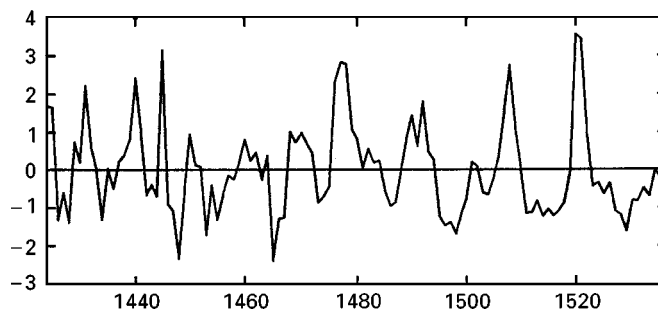


Fig. 5. Deviation from the long-run equilibrium (Equation 10)

framework.⁷ Consequently, efficient inference on the long-run relationship does not require knowledge of the way in which receipts are generated.

We can also use this marginal ECM for Δr_t to test the null hypothesis of Granger non-causality from e to r , by testing jointly the absence of all lagged Δe_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 significantly to 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 standard asymptotic distribution can be used 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 also has some interesting behavioural 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,

⁶ The inclusion of the dummy D1451 is necessary to have normality of the residuals.

⁷ 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 reparameterized. 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 that needs to be done is to estimate a (single equation structural) ECM for Δe_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 Δr_t using usual variable addition tests. 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 value of 2.737 well below the 5% critical value.

this could eventually explain the breakdown of the system. However, this will be implicitly analysed below.

V. THE STABILITY OF RECEIPT COLLECTION

If we are already able to point out some interesting facts about the relationship between receipt and expenditures in Nivelles over the sample period, the analysis up to now does not enable us to assess whether the model has truly isolated an invariant relation between the variables. In particular, if we want to be sure that the model estimated and formulated for the expenditures remains valid, even under various receipts regimes, we should establish 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 proposed recently by Engle and Hendry (1993), we will rely on an alternative way of assessing the empirical validity of the model. First, we must again complete the short-term model (Equation 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 section IV, a general-to-specific approach leads to the following specification:

$$\begin{aligned} \Delta r_t = & 0.84 \Delta f_t - 0.16 \Delta f_{t-2} + 0.06 \Delta r_{t-1} + 0.15 \Delta r_{t-2} \\ & (0.066) \quad (0.084) \quad (0.048) \quad (0.06) \\ & - 0.08 \Delta r_{t-3} + 2.24 D1437 \\ & (0.047) \quad (0.22) \end{aligned} \quad (11)$$

with

$$R^2 = 0.77 \quad DW = 1.85 \quad SER. COR. = 0.66$$

$$NORM = 18.22 \quad HET = 0.54$$

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.

The dynamic behaviour of the system is then investigated using dynamic simulations of Equations 9 and 11 independently (Figs. 6 and 7) and then jointly (Fig. 8). The underlying motivation is simple: if receipts are, effectively, super exogenous for the parameters of Equation 9, and not Granger-caused by expenditures, then the conditional model in Equation 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

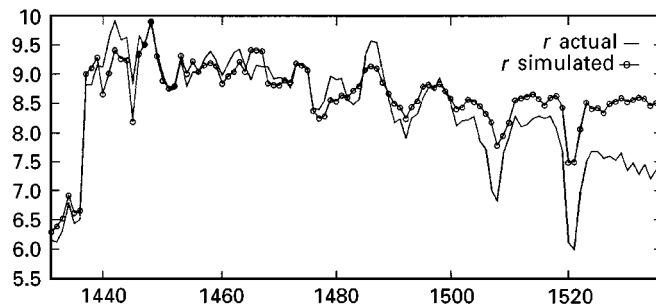


Fig. 6. Dynamic simulation of the marginal model

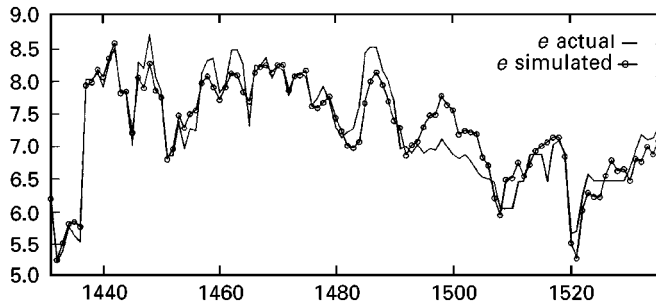


Fig. 7. Dynamic simulation of the conditional model

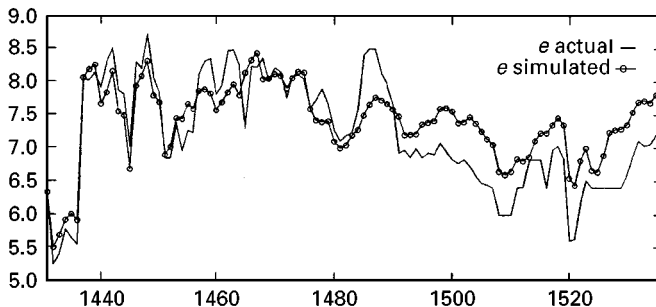


Fig. 8. Dynamic simulation of the full model

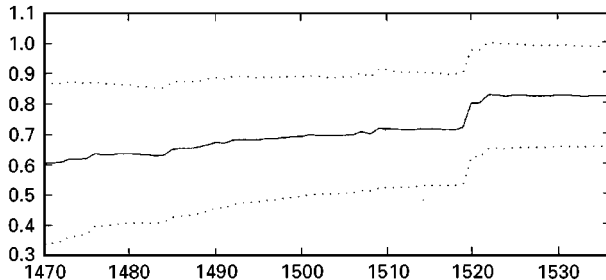


Fig. 9. Recursive short-run elasticity of receipts to fines

deteriorates the performance of the model for *e* in the joint simulation. This points out that, although the way in which 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, 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).

The interest of having derived a conditional ECM with valid conditioning variables is that the instability in receipt collection is limited to its source – in other words, 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 analysed. 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 following four elements.

A first element could be a change in the structure of fines. In general, fines are composed of fixed fines (called ‘faits mandés’) and pilgrimage fines. 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 proposition 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 as the offender’s choice, so that an actual journey will often replace a receipt flow.

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%–30% of fixed fines, while the lower fines (1–5 ‘oboles’) increased from 53% to 70%. Third, Rousseaux (1990) 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 1517–1536 reveals a change in the vocabulary: new terms related to corporal punishments are introduced (Lords’ justice becomes criminal justice). This change in vocabulary points to a change in the nature of penalties, implying that the proportion of monetary sanctions decreased during the 16th century, as illustrated in Table 9. Table 9 is also consistent with the fact that the emerging modern judicial system requires more formal proof of the offences, leading to a more intensive use of torture and physical punishment (see Langbein, 1974 and 1977).

To summarize: The instability in the link between receipts and fines can be explained by the evolution of the structure

Table 7. *Structure of fines*

	1457–1500	1500–1537
Fixed fines	64%	52%
Pilgrimages	15%	28%
Others	21%	20%

Table 8. *Structure of fixed fines*

	1457–1500	1500–1537
60 ‘oboles’	2%	0%
20 ‘oboles’	44%	24%
10 ‘oboles’	1%	6%
4 ‘oboles’	49%	64%
2 ‘oboles’	3%	4%
1 ‘oboles’	1%	2%

Table 9. *Monetary versus corporal sanctions*

	15th century		16th century	
	Number	%	Number	%
Monetary sanctions	4636	97.6	979	76.4
Corporal sanctions	113	2.4	302	23.6

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 to physical sanctions.

VI. CONCLUSIONS

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

The 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 mediaeval 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 (for example, 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 mediaeval, urban-based, tax-oriented system into a modern, State-based, punishment-focused system.

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