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Denise R. Osborn

EUI Working Paper RSC No. 95/15

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CRIME AND THE UK ECONOMY

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ABSTRACT

This paper investigates the quarterly time series for the number of notifiable offences recorded by the police in England and Wales over the period 1975 to 1993. The analysis concentrates on two major categories, namely total property and total personal crimes. After an investigation of the univariate characteristics, including seasonality and unit roots, the relationship between these crime categories and UK macroeconomic variables is considered. It is found that property crime may be cointegrated with real consumers' expenditure, but the long-run properties of personal crime cannot be explained by the macroeconomic aggregates. In the short-run, the major role is taken by changes in unemployment with both crime categories evidencing asymmetric responses to increases and decreases in unemployment.

1. INTRODUCTION

The nature of any relationship between crime and the economy is of wide interest. It is also a source of much debate. In a recent study, Pyle and Deadman (1994) review the extensive literature in the area, much of which investigates the potential link between crime and unemployment. This link is, indeed, the main focus of recent empirical studies by Reilly and Witt (1992) and Dickinson (1993), who use data relating to Scotland and to England and Wales, respectively. Both find evidence supporting the link.

Dickinson conceptually has a complex model, postulating an asymmetric relationship whereby "the increase in unemployment has a greater upward impact on criminal activity than does a decrease". Nevertheless, although he makes extensive use of graphs, his modelling is confined to simple linear regression using data expressed as percentage change series. He makes no attempt to test the asymmetry hypothesis in any formal manner. In concentrating on the crime/unemployment link, he also fails to consider the role of any other other potential explanatory variables. Reilly and Witt use regional data for Scotland and develop a pooled time series/cross section model, but they are also restricted in terms of the explanatory variables used.

The analysis of Pyle and Deadman (1994) builds on an important and extensive Home Office study by Field (1990). Field considers twelve categories of crime together with a number of economic indicators, demographic and weather variables. He finds that consumers' expenditure performs better than other economic variables, including unemployment, in explaining crime. Further, it appears that while property crime has an inverse relationship to the business cycle (consumption growth having a negative effect on the growth in such crimes), personal crime responds in a procyclical manner. Although takings Field's study as their starting point, Pyle and Deadman criticise him for conducting his analysis in terms of growth rates and, therefore, failing to model any long-run relationship between crime and economic variables. Despite their criticism of Field, the analysis of Pyle and Deadman is itself unsatisfactory in a number of respects. These are discussed in the next section.

The brief review above concentrates on recent UK time series studies of how crime relates to the economy, and the present paper constitutes another contribution in this sequence. Nevertheless, it needs to be emphasised that cross-section evidence is also available on this question. In this latter context, the most extensive UK data is that from the British Crime Surveys, which have to date been conducted in 1982, 1984, 1988, 1992 and 1994. These surveys ask approximately 10,000 individuals per sweep about property and personal crime experiences over the previous year. Extensive analysis of these has, however, failed to find any evidence of a positive influence from local unemployment rates on property crime (Osborn et al., 1992, Trickett et al., 1994). That is not to say that economic conditions are unimportant: indeed, Trickett et al. conclude that "richer people in poor areas suffer property crime particularly heavily".

The British Crime Surveys also throw some doubt on the accuracy of the police crime statistics used in time series studies (see, for example, Mayhew et al., 1994). Nevertheless, in common with the earlier studies of Field (1990), Dickenson (1993), Pyle and Deadman (1994), and others, police statistics are used for the analysis conducted here.

Section 2 considers the time series characteristics of recorded crime in England and Wales and also some aspects of the work of Pyle and Deadman. One important feature of our analysis, which has been almost completely overlooked in the previous literature, is the nature of seasonality in crime. In Section 3 we turn to the nature of any long-run relationship between the major crime aggregates and macroeconomic variables. Our results here confirm those of Field in finding consumers' expenditure to be the key macroeconomic indicator for explaining crime. The following section then examines the short-run: in particular, we examine evidence for an asymmetric relationship between movements in unemployment and those in crime. Conclusions in Section 5 complete the paper.

2. THE CRIME DATA AND ITS CHARACTERISTICS

Our data is the number of notifiable offences recorded by the police in England and Wales. This is available quarterly from 1975 onwards; our sample period extends to the end of 1993. Eight major offence categories are available: violence against the person, sexual offences, robbery, burglary, theft and handling of stolen goods, fraud and forgery, criminal damages, and other notifiable offences. There have been some definitional changes over our period, but only one of these has a substantial effect. That change involves criminal damage, where all offences have been recorded since the beginning of 1977, whereas damage of value £20 or less had previously been excluded.

In this analysis we consider four categories of crime, namely burglary, theft, criminal damage and personal crime. The first three are, of course, forms of property crime and the aggregate of the three is also considered. The last is formed as the sum of violence against the person plus sexual offences. These have been combined as the numbers in each of the two personal crime categories are relatively small, while they have very similar seasonality and other characteristics. Our four categories together constitute, on average, 95 percent of all recorded crimes. It may be noted that robbery, which cannot be comfortably categorised as a purely property or a purely personal crime, is excluded. As robbery accounts for less than 1 percent of all crimes, its omission is relatively unimportant in terms of the aggregates used in this analysis.

The crime data are shown in Figures 1 and 2 (pages 39-40). All three categories of property crime are included in the first graph, while personal crime and total property crime are shown (using different scales) in the second. As Figure 1 illustrates, theft is the most important property crime category, being larger than the aggregate of burglary and criminal damage in terms of the number of offences. Overall crime is, however, dominated by property ones with approximately 90 percent of all recorded offences.

It is notable that burglary and theft exhibit two periods of decrease, namely 1978/79 and 1988/89, with criminal damage also showing a slight decline in the latter period. On the other hand, personal crime (except for seasonal effects) increases almost continuously throughout the period. The step increase evident for criminal damage at the beginning of 1977 is due to the definitional change mentioned earlier.

Table 1 (pag.34) shows some of the important characteristics of crime in each of our categories, together with that for all recorded crime. The analysis of this table, and all subsequent analysis, is carried out after taking logarithms. The first part of Table 1 gives the average growth, while further statistics show the importance of seasonality. Finally, the table looks at the nature of the nonstationarity in this crime data; we return to this below.

The overall mean and the seasonal patterns shown in Table 1 have been computed from the regression

$$\Delta Y_t = \alpha_0 + \alpha_1(D_{1t} - D_{4t}) + \alpha_2(D_{2t} - D_{4t}) + \alpha_3(D_{3t} - D_{4t}) + u_t \qquad (1)$$

where ΔY_t is the first difference of the logarithm of the series, D_{it} is a dummy variable for quarter i (i=1,2,3,4) and u_t is a disturbance. The reported overall mean is the estimate of α_0, while the seasonal patterns for the first three quarters are the estimates of α_i for i=1, 2 and 3 respectively. For quarter 4, the restriction $\Sigma\alpha_i = 0$ yields the estimated seasonal pattern. Note also that the coefficient estimates from equation 1 have been scaled by 100 to to give the percentage values reported. Finally, SEE is the residual standard error (expressed as a percentage) and R^2 is the conventional coefficient of determination from equation (1). The form of (1) is identical to that in Osborn (1990) to enable seasonality in these crime variables to be compared with that of major UK macroeconomic aggregates.

Although the numbers in all categories of crime examined here have increased over the period, the fastest growth has been in criminal damage. Burglary and personal crime also exceed the growth rate for the total. Theft, which constitutes about half of all recorded offences, has grown on average more slowly than total crime. It is also notable from this first part of Table 1 that, overall, personal crime has increased at a faster rate than property crime.

The importance of seasonality in crime is, perhaps, surprising, but the extent is comparable to that for major real economic variables. Quarter to quarter movements are dominated by seasonality in the sense that the R^2 value for each category is at least a half. Indeed, the seasonal dummy variables in (1) explain 86 percent of the movements in personal crime; this figure is very similar to that for real consumption in the UK (Osborn, 1990, Table 3). In contrast to consumption, however, personal crime peaks in the summer months. This may, to some extent, be directly associated with seasonal climatic changes, in particular temperature and hours of daylight, making activities outside the home more attractive. It is, however, notable from Figure 2 that seasonality in personal crime does not appear to be constant over time: indeed, prior to the rapid increase experienced in the late 1980s, it seems that seasonality in personal crime was decreasing. It is also worth noting that Field (1990) finds beer consumption to be the most influential contemporaneous variable in explaining annual data on violence. This explanation cannot extend to the quarterly pattern, however, because beer consumption peaks in the Christmas quarter.

Economic theories of criminal behaviour, such as Becker (1968), usually view crime as deriving from utility maximisation. In this context it is notable that burglary, criminal damage and crime overall peak in the fourth quarter of the year, which is also the annual peak in many macroeconomic variables, but especially consumption (Barsky and Miron, 1989, Osborn, 1990). While it is not to deny that climatic seasonality, including hours of daylight, may play a role, the reduced criminal activity in the first quarter compared with the fourth suggests that criminals have seasonal utility functions similar to those of other consumers (Osborn, 1988). Theft and criminal damage are similar to each other in their seasonal patterns, with peaks in the second and fourth quarters. It is possible that a separation of vehicle from other thefts would clarify the pattern for that category, since seasonal patterns in thefts for material gain and thefts for enjoyment may exhibit different characteristics.

The third aspect of Table 1 is the nature of nonstationarity in recorded crime. In their study of annual postwar crime data for England and Wales, Pyle and Deadman (1994) find that the crime series are integrated of order two, or I(2). It is, however, widely accepted that real macoeconomic variables are I(1). As a consequence, Pyle and Deadman conclude that the levels of crime and economic activity cannot be cointegrated (Banerjee *et al.*, 1993), so that the level of the economy cannot explain the long-run level of crime. Therefore, they model the *change* in crime as a function of the *level* of the relevant explanatory economic variable. The consequence is that crime would continue to grow indefinitely in the long-run even if the level of economic activity is static. If true, this has profound implications.

There are, however, some difficulties with Pyle and Deadman's implementation of the unit root tests. To test the null hypothesis of an I(2) process against an I(1) alternative, a constant needs to be included in the augmented Dickey-Fuller test regression so that the critical values are not sensitive to the "starting value" of ΔY (Banerjee *et al.*, 1993, pp.104-105). At least for their major categories of burglary and theft, acceptance of I(2) by Pyle and Deadman is not entirely convincing when their results with a constant are examined. In any case, their use of an order 1 for augmentation is entirely arbitrary. Further, their error-correction mechanisms (ECMs) indicate that the dependent variable may be overdifferenced since the estimates of the error-correction coefficients are always close to minus one, which can be interpreted as the model attempting to reduce the order of differencing for the relevant crime variable[1]. In any case, the ECMs appear to be subject to dynamic misspecification. The only diagnostic check reported is the Durbin-Watson statistic, which (although biased towards failing to indicate the presence of autocorrelation in such a context) suggests the presence of positive residual autocorrelation in the annual residuals.

Except for the marginal (at 5 percent) test statistic for criminal damage, Table 1 gives no evidence that any of the crime series are I(2) over our sample period. Our estimates are based on a regression including an intercept and three

[1] Written in terms of a single explanatory variable for simplicity, their ECM is of the form $\Delta^2 Y_t = \beta_1 \Delta X_t + \beta_2(\Delta Y_t - \gamma_0 - \gamma_1 X_t) + \varepsilon_t$. With $\beta_2 = -1$ and $\gamma_1 = 0$, this becomes a model for ΔY in terms of ΔX.

seasonal dummy variables. Here, and in later analyses, the order of augmentation is chosen[2] to ensure satisfactory residual autocorrelation properties to order 4.

Continuing to "test down", we also examine I(1) versus I(0), with this latter test regression including a trend to give invariance to the value of a nonzero drift (Banerjee *et al.*, 1993, pp104-105). With test statistics close to the 5 percent critical values, these latter results indicate that the property crime categories and total crime may be more adequately described as trend rather than difference stationary. It is, however, also the case that Agiakloglou and Newbold (1992) find that a data-dependent augmentation order for the Dickey-Fuller test in the presence of a moving average component results in true significance values substantially exceeding the nominal ones. This, then, throws some doubt on results which are marginal in relation to the nominal 5 percent critical values and leaves the issue of difference versus trend stationarity unresolved for those categories. On the other hand, the I(1) hypothesis is clearly acceptable for personal crime. At least for the present, we proceed assuming difference stationarity, but we will revisit this briefly later.

[2] All regressions commenced with four lagged values of the dependent variable added. This order of augmentation was decreased or increased as indicated by the significance of estimated coefficients and the F-test version of the autocorrelation test.

Earlier discussion of Table 1 noted the importance of seasonality in the crime series. Therefore, the final part of Table 1 examines seasonality in the context of unit roots using an identical approach to that adopted in Osborn (1990). The initial hypothesis is that first and seasonal differencing are required, which is effectively a seasonal version of the I(2) hypothesis. Using an F-type statistic, the I(2) hypothesis is once again very clearly rejected. When seasonal versus conventional first differencing is considered, the results for burglary, theft, total property crime and all crimes indicate that first differencing only is required. On the other hand, the test is unable to distinguish clearly between first and seasonal differencing for criminal damage and total personal crime.

3. CRIME AND ECONOMIC VARIABLES IN THE LONG-RUN

Figure 3 (pag. 41) shows our major categories of property and personal crimes, together with two key macroeconomic variables, namely unemployment and total real consumers' expenditure. As all series are analysed in logarithmic form, logs have been taken for the graphs in Figure 3. As with the analysis of Table 1, no seasonal adjustment has been applied to the crime data. For unemployment, however, no consistent seasonally unadjusted data are available over the period, so that the series shown in Figure 3 is the total number unemployed on a seasonally adjusted basis. Consumers' expenditure is readily available unadjusted and it exhibits rich seasonal characteristics (Osborn, 1988). Nevertheless, for the analysis of this section we wish to emphasise the long-run movements, so that it is graphed in the form of a moving annual average of log real consumption in the current and immediately preceding three quarters.

As noted in Section 1, the effect of unemployment on crime has been of considerable interest in the literature of criminology. However, Figure 3 indicates that a long-run *linear* relationship with unemployment cannot be used as an explanation of the rise in crime over this period. This is obvious in that roughly comparable levels of unemployment in 1986 and 1993 are associated with very different crime levels. Indeed, total crime is about a third higher at the later date when the peak of unemployment is actually lower. Further, as Field (1990) notes over his longer period, there is also a problem in relating periods of increasing crime with increasing unemployment, because the upturn in crime typically pre-dates that of unemployment[3]. Thus, we conclude that there is no linear cointegration between unemployment and crime.

[3] Dickinson (1993) argues that youth, not total, unemployment is the relevant measure for explaining crime and that youth unemployment has different turning points. The measure he uses is not, however, available quarterly.

Field stresses the role of real consumption in explaining short-run changes in crime. Since he models using differenced data, he does not explicitly address the issue of any long-run relationship. Nevertheless, there are plausible reasons why such a long-run relationship might exist. Crime, whether property or personal, yields utility to the offender. It is then reasonable to assume that offenders will aspire to similar levels of consumption as those attained by the population as a whole. This is implies a long-run positive relationship. Any such relationship is, however, likely to be modified by many other factors, including the ease with which legal employment can be obtained.

The analysis here of crime and the economy is not to imply that only economic factors are important for crime in the long-run. It seems obvious that social and criminological influences will also be important. Nevertheless, these factors are not only complex, but they are also closely interrelated. Thus, for example, the changes which have resulted in decreasing social control on individual behaviour have been associated with a period of long-term growth in the economy. Here we take the position that although it is conceptually impossible to completely separate the many long-run influences on crime, they may be proxied by macroeconomic aggregates. As a single macroeconomic variable, we concentrate on real consumers' expenditure as capturing the long-run aspirations of criminals.

As indicated by Figure 3, any long-run relationship of crime with consumption will be positive. However, the rapidly rising crime rate of the late 1980s and early 1990s then represents a substantial (and prolonged) deviation from the long-run, because consumption is essentially static at this time. The issue of cointegration is addressed in Table 2 using the first-stage of the Engle-Granger (1987) two-step method. Given the relatively short time series available, this is preferred to the vector autoregressive approach of Johansen (1988). In addition to showing the results of a regression of each crime category on consumption, we also include results using real gross domestic product (GDP). As the Central Statistical Office has recently ceased producing an unadjusted GDP series, that series is also seasonally adjusted.

Consumption, as used in Table 2 (pag. 35) and Figure 3, is expressed as the average of the annual moving sum. This annual moving sum removes the nonstationary seasonal unit roots exhibited by the unadjusted series (Osborn, 1990), leaving the non-seasonal long-run component of interest here (Hylleberg et al., 1990). The results were also checked by using untransformed quarterly consumption, but the annual average always performs better. Also note that, to allow for the discontinuity in the definition of criminal damage, a dummy variable taking the value one in each quarter of 1975 and 1976 is included where appropriate. All the cointegrating test regressions include seasonal dummy variables.

For the larger aggregates, namely total property, total personal and all crimes, the evidence of the augmented Dickey-Fuller statistic indicates that a long-run relationship may, indeed, be present between each of these variables and real consumption. Our results confirm that consumption explains long-run crime levels better than GDP. The cointegration results for personal crime are not, however, convincing when examined further. The Dickey-Fuller test regression without augmentation only just passes the residual autocorrelation test at a 5 percent level (the p-value is .078). With an augmentation of 4 lags the residual diagnostics are more satisfactory, but the ADF statistic becomes -2.11. There is, in any case, the possibility that personal crime contains seasonal unit roots. If the annual average value is used as the dependent variable to remove such roots, the ADF test (using 4 lags) is -1.08. Finally, cointegration of both property and personal crime with consumption would imply that the two series were themselves cointegrated, but a direct test yielded no evidence that such cointegration exists.

Although we noted that unemployment could not be cointegrated with crime in a linear bivariate context, that did not preclude the possibility that unemployment plays some role in conjunction with a variable such as consumption. This is investigated in the second part of Table 2. It is striking, however, that the addition of unemployment results (in almost every case) in weaker evidence for cointegration.

In the next section we move from the long-run to a complete specification of the dynamic relationship between crime and the economy. This modelling is pursued at the intermediate aggregation level of total property and total personal crimes. Because of the inherently different nature of the two types of crimes, it seems meaningful to keep them separate. Further, we have just concluded that personal crime may not maintain a long-run (linear) relationship with the economy. For property crime, we use consumption as the sole long-run explanatory variable. The estimated elasticity exceeding unity (approximately 1.5 in Table 2) implies that criminals aspire to more than match aggregate increases in consumption. For personal crime, we abandon long-run modelling and consider the extent to which short-run changes can be explained[4].

[4] Further investigation of a long-run relationship for personal crime and consumption was undertaken in the context of a single stage estimation of a short/long run model, but no satisfactory results were obtained.

4. MODELLING SHORT-RUN DYNAMICS

Prior to undertaking any short-run modelling, one adjustment was made to the property crime series to allow for the definitional change which occurred at the beginning of 1977. The adjustment used the dummy variable included for 1975 and 1976 in the cointegrating regression of property crime with average consumption, and the value of the dummy variable coefficient was added to each of the (log) property crime values for this period.

To induce stationarity, first differences are applied to property crime, but seasonal (annual) differences are taken for personal crime. There is no question of taking seasonal differences for the former according to the unit root tests of Table 1, but these tests are inconclusive on whether first or seasonal differences are appropriate for personal crime. In the context of the models in this section, however, annual differences resulted in more satisfactory models.

The more satisfactory nature of seasonal differences for personal crime can be explained by Figure 3. There it appears that the seasonal movements in this variable decline in magnitude over time, indicating possible nonstationarity in these movements. Although the earlier unit root tests were not decisive, this visual evidence, and the more satisfactory short-run models found, led to the adoption of seasonal difference in personal crime for short-run modelling.

The (first or seasonal) differenced variables employed in the short-run modelling are graphed in Figure 4 (pag. 42).

The Granger Representation Theorem (Engle and Granger, 1987) states that if a cointegrating relationship exists, then the short-run dynamic relationship can be written in error-correction form. Here we start with the Engle-Granger two-step method and include the lagged residual from the property crime/consumption cointegrating equation (Table 2) in a dynamic model for differenced property crime. The initial specification was general, containing four lags of total property and total personal crime, together with the contemporaneous and four lagged differences of real consumption and unemployment. First differences were used for unemployment, whereas seasonal (annual) ones were employed for consumption as indicated by the seasonal unit roots in the series (Osborn, 1990).

Due to the interest in cyclical asymmetry, as emphasised by Dickinson (1993), two further variables were included, namely a dummy variable taking the value unity when unemployment is increasing and another dummy which is unity when consumption is increasing, the latter being compared to the corresponding quarter of the previous year. A constant and three seasonal dummies were also in the initial model for property crime.

With two exceptions, an identical initial specification was used for personal crime. The exceptions are that no lagged cointegrating residual was included and, since seasonal differences were used, seasonal dummy variables were redundant. When the property crime equation was estimated, however, some residual autrocorrelation was indicated and an eight period lag of the dependent variable was included. The coefficient estimates, obtained by ordinary least squares applied to these general models[5], are summarised in Table 3 (pag. 36).

For property crime, the error-correction term is individually significant well beyond the 5 percent level, even in this general over-parameterised specification. Lagged dependent variable values are important, but it is possibly more surprising that lagged personal crime contains significant information for property crime changes in the short-run. The effect here is negative, in line with the different directions of short-run movement noted by Field (1990). Overall,

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[5] The diagnostic tests reported are conventional. The significance levels are obtained using the F-test variants of a test for autocorrelation to order 4, the addition of the squared fitted value (linearity) and a regression of the squared residual on the squared fitted value. The normality test is a χ^2 test of skewness and kurtosis.

changes in consumption and unemployment appear to have little or no role in determining short-run property crime. An asymmetric effect for unemployment is, however, given credence by its significant positive coefficient when included as a simple sign dummy variable. On the other hand, increasing consumption seems to be unimportant as an explanation of property crime changes. The diagnostic tests do not throw up any problems with this equation. The results of the personal crime equation are discussed below.

Before leaving the property crime equation of Table 3, one further remark is in order. It will be recalled that Table 1 left the question of trend versus difference stationarity unanswered for property crime. The issue was examined again in the context of the general model in Table 3 by comparing two specifications of the long-run, while maintaining the presence of short-run dynamics. In one the lagged cointegrating residual was replaced by a trend and the one-period lag of the level of property crime, while in the other lagged average consumption and lagged property crime levels were used. The two sets of results were very similar, with R^2 values of .809 for the former and .808 for the latter. It remains true, therefore, that the data is unable to distinguish clearly between these two long-run specifications. On a priori grounds, we prefer the explanation provided by consumption.

Beginning from the results of the general specification for property crime in Table 3, the model was refined. The strategy was to drop insignificant lagged (property and personal) crime terms and then insignificant consumption and unemployment terms. The failure to find any role for changes in consumption and unemployment was confirmed even when only one lag of each was included. Finally, the role of the two asymmetry dummy variables was investigated. It was found that the signficance of the unemployment one improved slightly when lagged by one quarter, with the lag 8 on the dependent variable then becoming insignificant. The consumption dummy was never close to a 5 percent signficance level for any specification. The final equation obtained by this process is shown in the first column of Table 4 (pag. 37). Once again, results are shown in terms of each estimated coefficient and its (two-sided) significance level.

In relation to the general specification of Table 3, the short-run dynamic equation for property crime in Table 4 holds no surprises. It is clear that recent past movements in personal crime retain a powerful role in explaining current property crime, indicating a possible substitution over time of property crime for personal crime. The annual lag for property crime itself suggests that the behaviour of criminals may include an element of seasonal habit-persistence, as does the action of other consumers (Osborn, 1988). The only short-run economic variable in this final specification is the lagged dummy indicating periods of increasing unemployment. As seen in Figure 3 or 4, swings in unemployment are smooth so this variable maintains a zero or one value for relatively long periods. There is, indeed, evidence to support Dickinson's (1993) asymmetry hypothesis here, with crime increasing faster when unemployment grows than it declines in the opposite case. The effect captured here is, however, purely a short-run one.

The model was checked in a number of ways. To investigate whether asymmetry was pervasive, all the coefficients of the model were allowed to take different values when lagged unemployment changes were positive and when they were negative. No role could be found, however, beyond that of the simple switch dummy. The two-stage estimation was also collapsed into a single equation by explicitly including levels of lagged property crime and lagged average consumption; the results are shown in the second column of Table 4. The effects were relatively slight, although the implied long-run elasticity of property crime with respect to consumption rises from 1.5 in Table 2 to 1.7 here. Further, the equation was estimated to the end of 1985, with the final 32 observations used to check structural stability. This may be seen as a tough test since the date of the "break" was chosen to pre-date the latest increase in unemployment.

Nevertheless, the Chow test of structural stability is comfortably passed. Although there is some evidence of nonlinearity in the equation estimated over this shorter period, the coefficient estimates themselves are quite robust.

The final column of Table 4 records the results of a joint estimation of the property and personal crime equations. We will return to that after considering the investigations of personal crime in a single equation context.

As indicated by its diagnostics, the general specification in the second column of Table 3 for property crime is unsatisfactory. The linearity and heteroscedasticity checks are not independent, but the problem seems to be that changes in personal crime are not a linear function of the variables employed here. Further investigation was undertaken and the key appeared to be an interaction of the unemployment change dummy variable with unemployment itself. Indeed, satisfactory diagnostics were obtained by the addition of a separate variable which is zero when unemployment is decreasing, but which takes the value of the unemployment change when it is increasing[6]. This was included with the same lags as unemployment itself; both unemployment changes and positive unemployment changes are highly significant, as seen in the final

[6] This variable was created by multiplying the increasing unemployment dummy variable by the change in unemployment.

column of Table 3. Indeed, these results indicate that *only* unemployment changes are important in the explanation of personal crime movements. The success of this specification does, however, indicate that asymmetric responses to unemployment are far more pervasive for personal than for property crime.

Now starting from this new general specification, an analogous approach to that for property crime was taken to specifying a more parsiminous personal crime equation. In this context, an annual lag of the dependent variable was required; the coefficient for this variable was also found to differ significantly with the sign of the change in unemployment. No other variable was, however, found to be important. The final personal crime equation is shown in the first column of Table 5 (pag. 38). Here the specification of lagged unemployment differs slightly[7] from that in the final column of Table 3.

Before looking further at Table 5, it is worth examining further the relationship between changes in personal crime and changes in unemployment as revealed by Figure 4. The inverse relationship between the two variables appears obvious, but it is particularly marked in times of decreasing unemployment.

[7] In the latter the sign of the appropriate lagged unemployment variable was used in defining the positive change in unemployment variable, whereas Table 5 is a switching model where the switch is defined in terms of the sign of current unemployment changes. These differ only when the sign changes between the lagged and current value.

Especially notable is the latter part of the 1980s when decreasing unemployment seems to be related to large positive changes in personal crime. It is, however, less clear that moderate increases in unemployment have any relationship at all with patterns in these crimes.

That message from Figure 4 lies behind the results of Table 5. The coefficient estimates of the first column imply that increasing unemployment has much less effect than does decreasing unemployment. Further, the lagged crime variable has less influence when unemployment is rising. Despite the individual lack of significance of the three variables (other than the constant) in the increasing unemployment regime relationship, the variables are jointly highly significant (significance level .006).

As with property crime, other checks were carried out. Not surprisingly, a test for equal slopes for the three explanatory variables (lagged personal crime, current and lagged unemployment) across increasing and decreasing unemployment regimes clearly rejected equality. Further, when estimated to the end of 1985, the structural stability test was comfortably passed. This is quite remarkable, since it implies that the equation estimated to 1985 is compatible with the steep increase in personal crime experienced in the late 1980s; see Figure 3. Indeed, it is worth noting that the coefficients in the first and second columns of Table 5 are very similar. The only exception is that the role of lagged personal crime in the estimates over the whole period seems to derive principally from 1986 onwards.

The final econometric investigation undertaken here is a system estimation of the two crime equations. This exploits the fact that the equations form a seemingly unrelated system and, if there is correlation in the disturbances across the two equations, then estimation exploiting this will yield increased efficiency. In practice, although a correlation of .21 was found between the two residual series, the coefficient estimates differ little from those obtained using ordinary least squares. Nevertheless, the results are shown in the final columns of Tables 4 and 5. A χ^2 test for vector residual autocorrelation to lag 4 yields a satisfactory significance level of .38.

5. CONCLUSIONS

This paper set out to examine the relationship between crime and UK macroeconomic aggregates in the context of quarterly crime data. Although the sample period covers less than twenty years, potentially important relationships have been uncovered. Possibly the most important of these is the link between consumption and crime. This was first emphasised by Field (1990) in the context of the annual growth in crime, but here the effect of consumption appears to be even more important. Although not pursued, the seasonality analysis in Section 2 suggests that quarterly patterns in property crime may be linked to a seasonal utility function underlying real consumers' expenditure. For property crime once again, consumption appears to provide a long-run explanation of the increase in crime. Short-run movements are, however, modelled using the dynamics of crime and a dummy variable for periods of increasing unemployment.

Personal crime is less obviously linked to consumption. Indeed, we could find no long-run economic explanation for personal crime, while in the short-run evidence was found of asymmetric responses to increases and decreases in unemployment. Such asymmetry would, of course, explain the failure to find a linear long-run relationship. If personal crime increases steeply with decreases in unemployment, but responds little when unemployment rises, then a "ratchet" effect will result which would not be captured by a linear long-run specification. Although we do not find a role for consumption, the results can be interpreted in a similar way to those of Field (1990, pp.35-36). He argues that in periods of increasing consumption, individuals spend more time outside the home and this results in more victimisation. Here the same argument can be applied when unemployment is decreasing. An extra element is, however, required for the observed "ratchet" effect. This is compatible with people becoming used to their way of life and so not readily adjusting back to a less active social life when unemployment rises.

Another interesting aspect of the results is that past changes in personal crime contain explanatory information for property crime, but the reverse does not apply. The negative effect appears to imply substitution over time between personal and property crime. Although movements in unemployment influenced short-run property crime only through a lagged dummy variable indicating periods of increasing unemployment, it should be borne in mind that further effects over time are indicated through lagged personal crime.

Our property crime model supports Dickinson's (1993) hypothesis that increases in unemployment have a "greater upward impact on criminal activity than does a decrease". Nevertheless, the effect we capture is a short-run one. It is also worth recalling that property crime trends upwards in a manner which makes it difficult to distinguish whether the series is difference or trend stationary. It is possible that the explanation of long-run property crime movements are more complex than attempted here, with an effect from consumption possibly combined with an asymmetric ratchet effect from unemployment.

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Table 1. The Characteristics of Crime

Statistic	Crime Type					
	Burglary	Theft	Criminal Damage[a]	Total Property[a]	Total Personal	All Crime[a]
Av. Growth[b]	1.30	0.97	1.72	1.11	1.29	1.10
Seasonal Patterns[b]: Deviation from Overall Mean						
Quarter 1	2.84	-4.92	-3.21	-2.70	-12.14	-3.24
Quarter 2	-5.91	4.53	5.60	1.89	14.42	2.35
Quarter 3	-6.79	-1.66	-8.54	-4.04	4.51	-3.42
Quarter 4	9.86	2.04	6.15	4.85	-6.80	4.30
SEE	4.91	3.54	5.19	3.68	4.18	3.51
R^2	.674	.518	.603	.500	.864	.500
Unit root tests[c]						
I(2) v I(1) (augmentation)	-7.47 (0)	-7.36 (0)	-3.53 (3)	-7.77 (0)	-8.41 (2)	-8.04 (0)
I(1) v I(0) (augmentation)	-3.11 (4)	-3.43 (4)	-3.08 (5)	-3.72 (4)	-1.36 (3)	-3.63 (4)
Seasonal unit root tests[d]						
H_0: $\Delta_1\Delta_4$ (augmentation)	5.83 (4)	21.50 (2)	19.25 (0)	28.51 (0)	28.42 (0)	16.32 (0)
H_a: Δ_4 only	-0.57	-0.56	-3.01	0.81	-3.58	0.69
H_a: Δ_1 only	-2.67	-4.82	-3.71	-5.28	-4.94	-5.22

Notes: a. Computed using data from 1977(1) only.

b. Figures (except for R^2) are expressed as percentages.

c. Approximate 5% and 1% critical values (Fuller, 1976) are: -2.89 and -3.51 for I(2) v I(1), -3.45 and -4.04 for I(1) v I(0).

d. Approximate 5% and 1% critical values (Osborn, 1990) are: 3.79 and 4.80 for overall test -2.11 and -2.82 for Δ_4 alternative, -3.75 and -4.35 for Δ_1 alternative.

Table 2. Testing the Long-run Relationship of Crime with Economic Variables

	Burglary		Theft		Criminal Damage[a]		Total Property[a]		Total Personal		All Crimes[a]	
Statistic	Cons.[b]	GDP	Cons.[b]	GDP	Cons.[b]	GDP	Cons.[b]	GDP	Cons.[c]	GDP	Cons.[b]	GDP
Coefficient	1.59	2.07	1.40	1.84	2.20	2.98	1.49	1.98	1.82	2.41	1.49	1.99
R^2	.730	.691	.836	.807	.968	.957	.864	.829	.969	.947	.871	.836
ADF[c] (augmentation)	-3.40 (4)	-3.17 (4)	-3.70 (4)	-2.78 (2)	-2.68 (4)	-2.91 (4)	-4.13 (4)	-3.88 (4)	-4.15 (0)	-1.39 (2)	-4.08 (4)	-3.76 (4)
Adding unemployment to cointegrating regression												
Cons./GDP coeff.	1.11	1.39	1.17	1.52	1.92	2.57	1.27	1.68	1.76	2.33	1.29	1.70
Unempl. coeff.	.320	.333	.152	.157	.277	.302	.210	.230	.039	.042	.196	.215
R^2	.875	.845	.885	.859	.986	.980	.916	.892	.972	.949	.918	.894
ADF[c] (augmentation)	-2.48 (4)	-1.91 (3)	-2.91 (2)	-2.67 (2)	-2.17 (4)	-2.17 (4)	-2.93 (4)	-2.49 (4)	-4.04 (0)	-1.44 (2)	-2.44 (2)	-1.63 (0)

Note: All regressions include an intercept and three seasonal dummy variables.
a. These regressions also include a dummy variable for 1975 and 1976.
b. Consumption is expressed as the average value over the year.
c. Approximate critical values (MacKinnon, 1991) are: -3.42 and -4.05 at 5 and 1 percent respectively for 1 regressor, -3.86 and -4.49 for 2 regressors.

Table 3. General Models of Short-Run Dynamics: Dependent Variables are $\Delta_1 prop_t$ and $\Delta_4 pers_t$

Variable	Total Property	Total Personal	Total Personal
Error-correction	-.214 (.000)		
Change in property crime Lags 1,2,3,4[a]	.868 (.013)	-.140 (.940)	.296 (.659)
Change in personal crime Lags 1,2,3,4	-.345 (.008)	-.032 (.371)	-.564 (.362)
Change in consumption Lags 0,1,2,3,4	-.095 (.938)	.409 (.588)	.166 (.479)
Change in unemployment Lags 0,1,2,3,4	-.248 (.494)	-.487 (.085)	-2.332 (.000)
Positive change in unemployment Lags 0,1,2,3,4			1.883 (.002)
Increasing unemployment dummy	.0304 (.022)	.0038 (.821)	.0332 (.058)
Increasing consumption dummy	-.0040 (.797)	-.0248 (.206)	-.0222 (.200)
R^2	.804	.507	.680
SEE	.029	.039	.034
Diagnostic tests: Significance levels			
Autocorrelation	.283	.521	.209
Linearity	.484	.001	.122
Normality	.348	.647	.718
Heteroscedasticity	.293	.015	.976

Note: The value shown is the sum of the individual coefficients, with the (two-sided) significance level for a joint test of zero coefficients given in parentheses.

a. Lag 8 is also included for property crime equation.

Table 4. Property Crime Equation
Dependent Variable is $\Delta_1 prop_t$

Variable	Second-stage Equation 1975-1993	One-stage Estimation 1975-1993	One-stage Estimation 1975-1985	System Estimation 1975-1993
EC_{t-1}	-.186 (.000)			
$cons_{t-1}$		.344 (.000)	.388 (.014)	.337 (.000)
$prop_{t-1}$		-.203 (.000)	-.200 (.010)	-.201 (.000)
$\Delta_1 prop_{t-4}$	.452 (.000)	.412 (.000)	.307 (.049)	.432 (.000)
$\Delta_4 pers_{t-1}$	-.322 (.000)	-.326 (.000)	-.354 (.011)	-.331 (.000)
$\Delta_4 pers_{t-2}$	.204 (.021)	.173 (.049)	.168 (.204)	.187 (.019)
$Dum(\Delta_1 unemp_{t-1} > 0)$	.0287 (.002)	.0346 (.000)	.0270 (.082)	.0342 (.000)
Constant	-.0220 (.045)	-1.101 (.001)	-1.637 (.104)	-1.052 (.000)
Quarter 2	.0248 (.025)	.0194 (.076)	.0478 (.015)	.0194 (.049)
Quarter 3	-.0070 (.448)	-.0094 (.301)	-.0009 (.947)	-.0081 (.333)
Quarter 4	.0396 (.002)	.0330 (.011)	.0497 (.013)	.0341 (.003)
R^2	.753	.767	.800	.767
SEE	.027	.027	.028	.027
Diagnostic tests: Significance levels				
Autocorrelation	.192	.103	.172	
Linearity	.341	.226	.026	
Normality	.193	.312	.686	
Heteroscedasticity	.800	.979	.914	
Chow test			.701	
Cyclically varying slopes	.747			

Table 5. Personal Crime Equation
Dependent Variable is $\Delta_4 pers_t$

Variable	Single Eqn. Estimation 1975-1993	Single Eqn. Estimation 1975-1985	System Estimation 1975-1993
Increasing unemployment regime			
$\Delta_4 pers_{t-4}$	-.164 (.201)	-.097 (.583)	-.152 (.193)
$\Delta_1 unemp_t$	-.196 (.375)	-.163 (.583)	-.158 (.436)
$\Delta_1 unemp_{t-1}$	-.201 (.339)	-.287 (.309)	-.200 (.301)
Constant	.0603 (.000)	.0558 (.000)	.0579 (.000)
Decreasing unemployment regime			
$\Delta_4 pers_{t-4}$	-.522 (.000)	-.143 (.736)	-.490 (.001)
$\Delta_1 unemp_t$	-1.479 (.002)	-2.000 (.166)	-1.528 (.000)
$\Delta_1 unemp_{t-1}$	-.855 (.024)	-.670 (.581)	-.777 (.023)
Constant	.0375 (.004)	.0223 (.638)	.0363 (.002)
R^2	.598	.370	.596
SEE	.032	.035	.032
Diagnostic tests: Significance levels			
Autocorrelation	.989	.957	
Linearity	.175	.635	
Normality	.800	.988	
Heteroscedasticity	.942	.541	
Chow test		.517	
Cyclically varying slopes	.000		

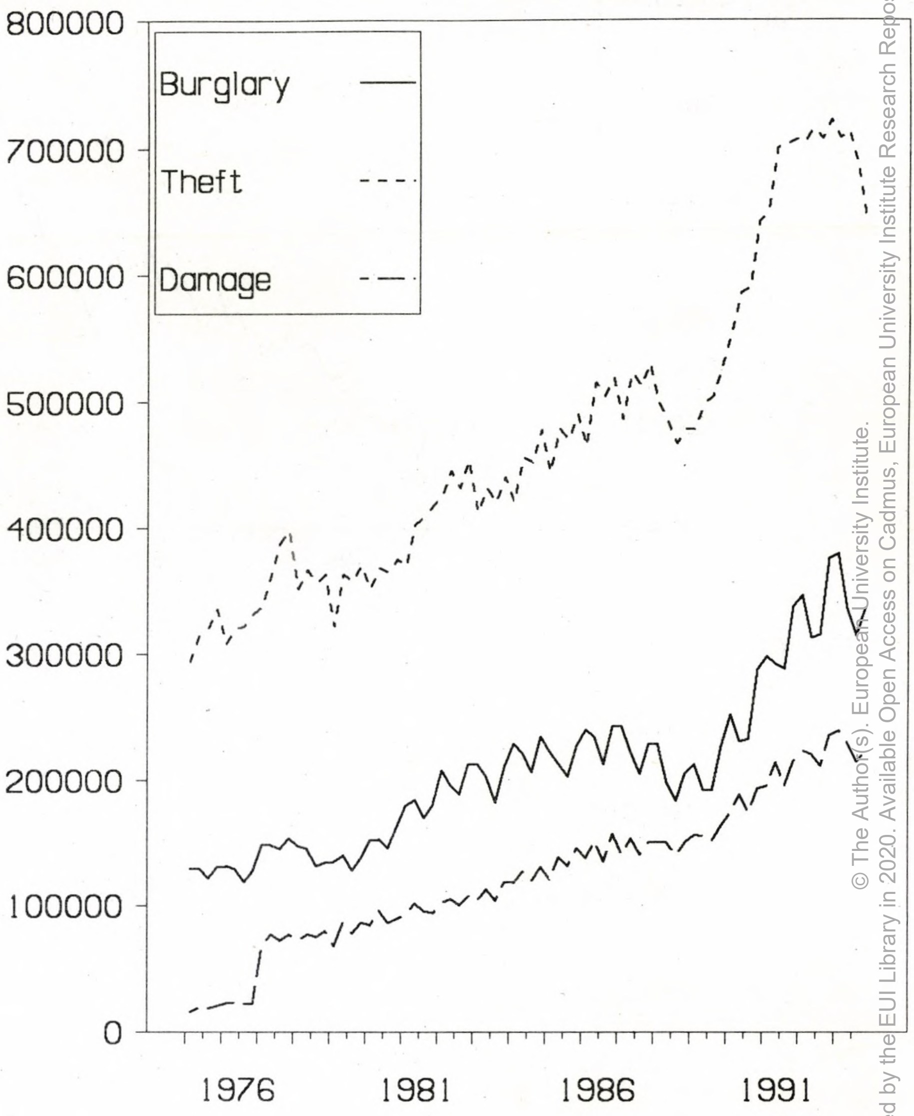

Figure 1. Quarterly series on the component categories of property crime.

(a)

(b)

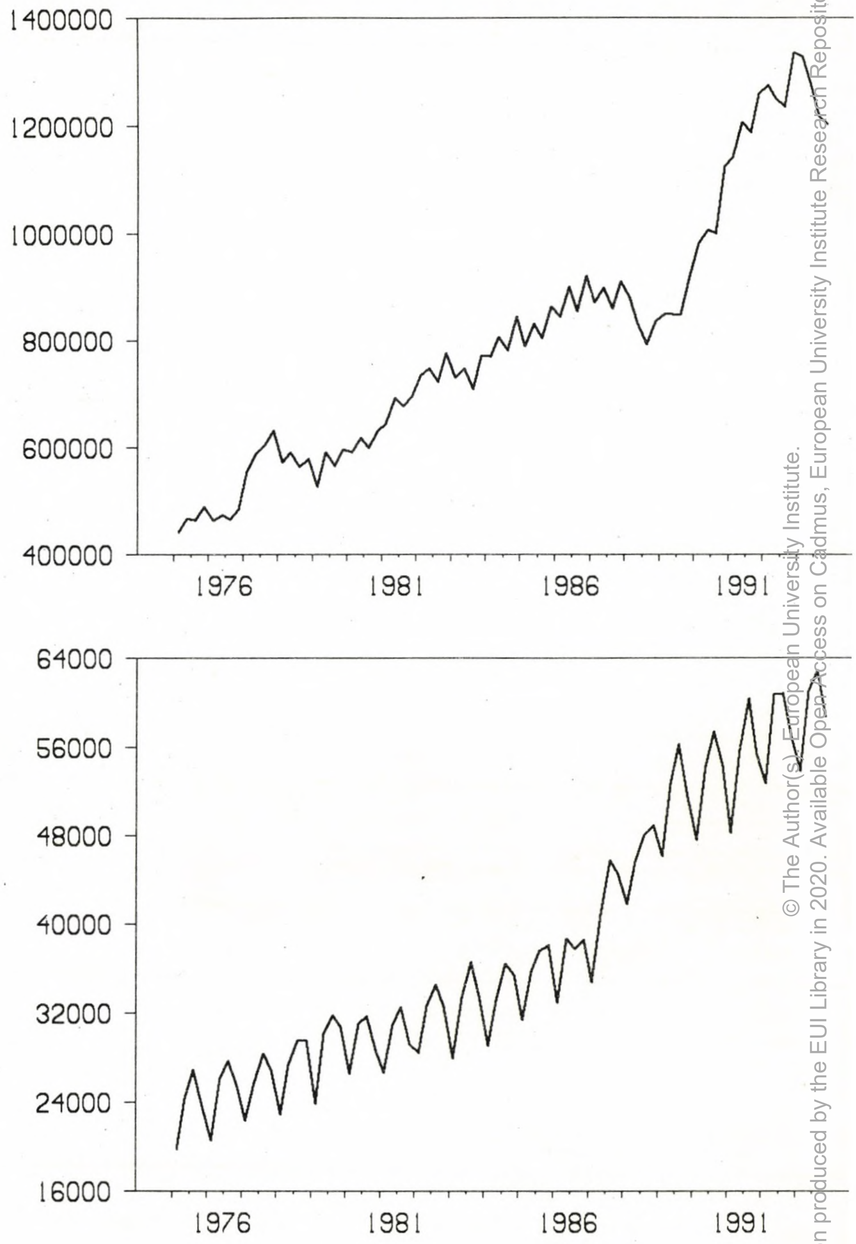

Figure 2. (a) Aggregate property crime and (b) Aggregate personal crime.

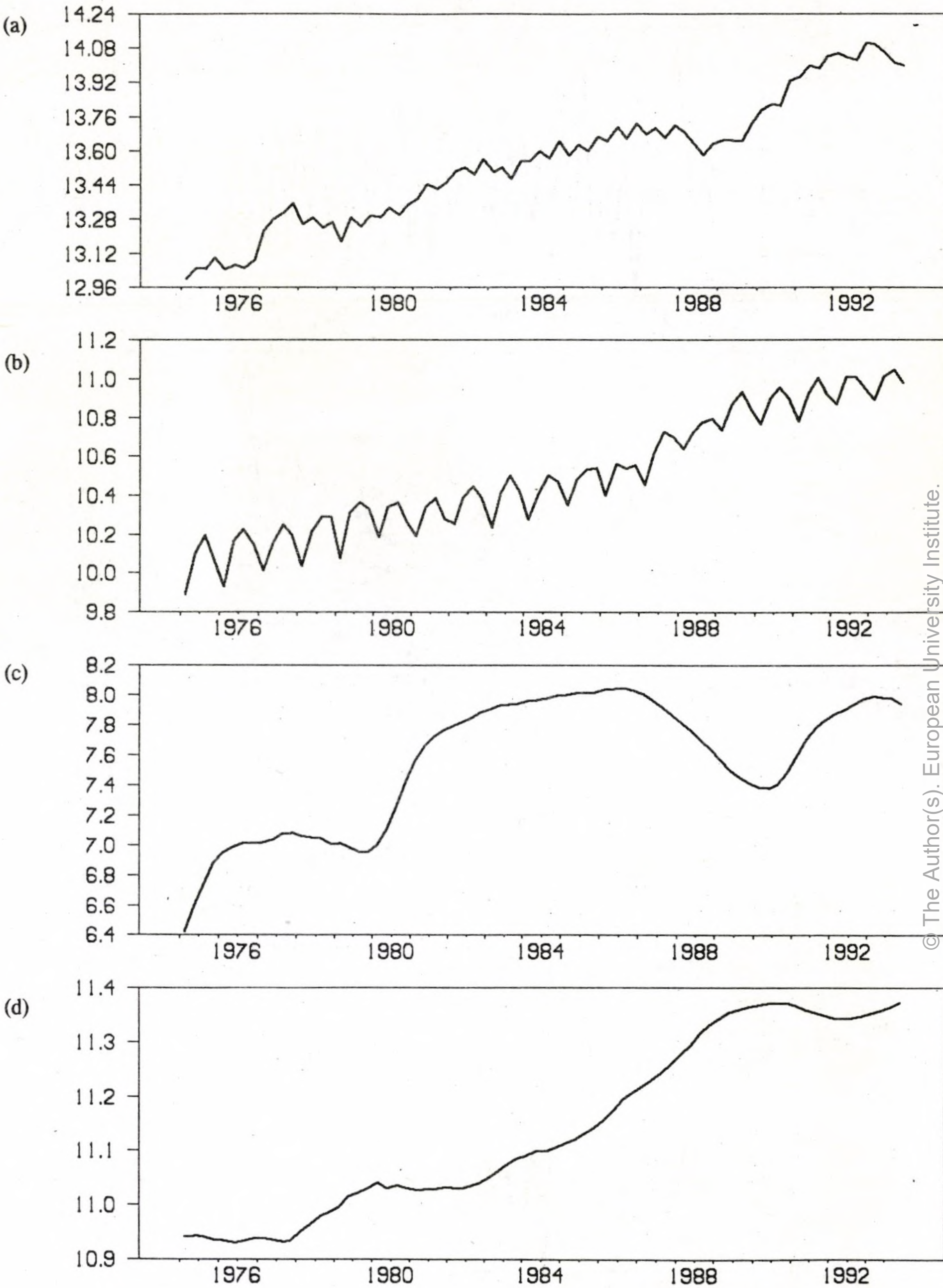

Figure 3. (a) Aggregate property crime, (b) Aggregate personal crime, (c) Unemployment and (d) Moving annual sum of real consumption; all expressed in logarithms.

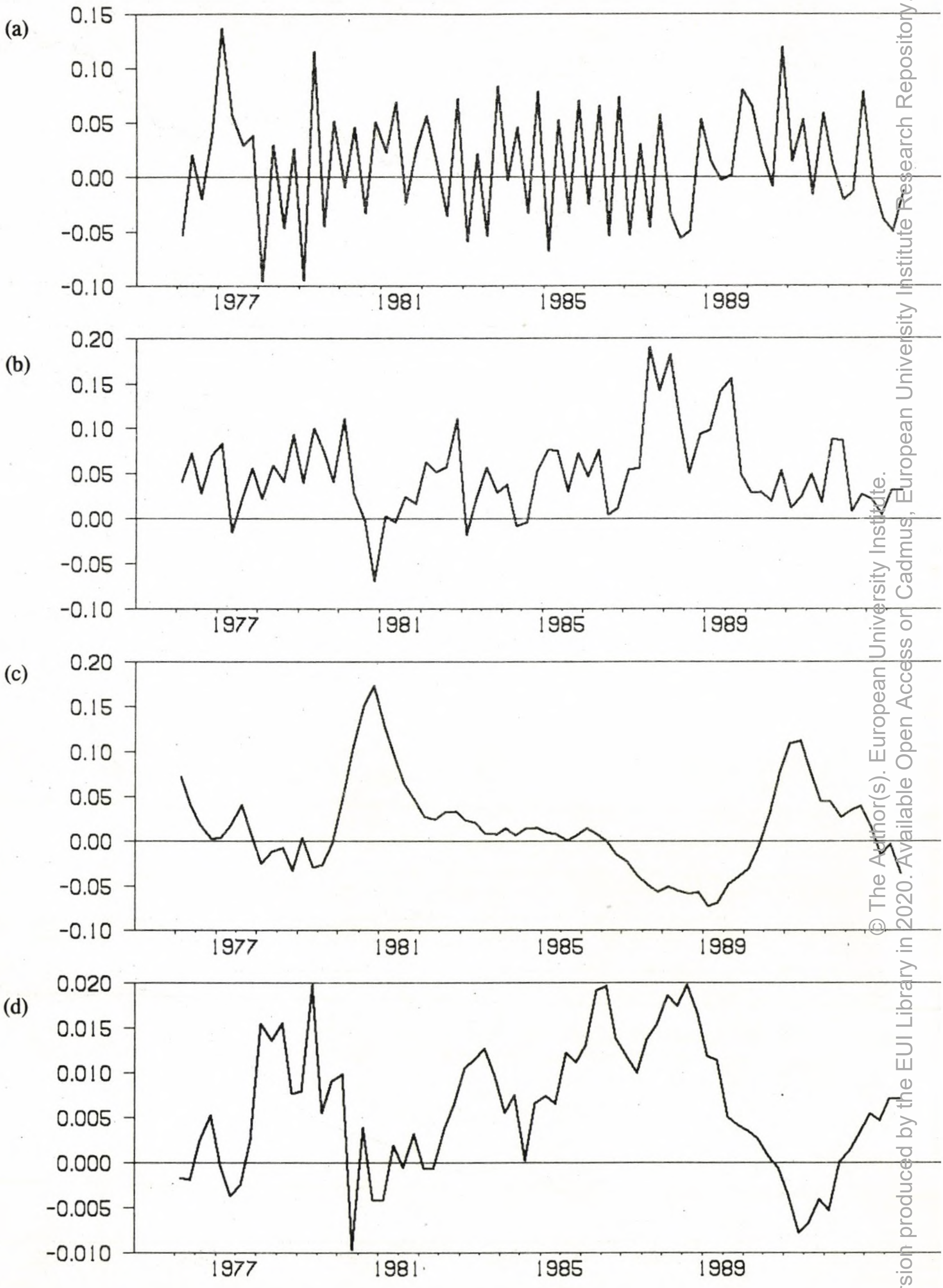

Figure 4. Differenced variables used in short-run modelling: (a) Aggregate property crime, (b) Aggregate personal crime, (c) Unemployment and (d) Real consumption.

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