Bank of England

Discussion Papers

Technical Series

No 48

Secondary market prices of LDC debt

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The object of this Technical Series of Discussion Papers is to give wider circulation to research work in the Bank, and to invite comment upon it; any comments should be sent to the authors at the address given below.

The authors would like to thank numerous Bank of England colleagues for their comments and suggestions and, in particular, Andrew Bailey, Peter Doyle and Colin Miles. All remaining errors are the responsibility of the authors. The views expressed are those of the authors and not necessarily those of the Bank of England.

Issued by the Economics Division, Bank of England, London, EC2R 8AH to which requests for individual copies and applications for mailing list facilities should be addressed; envelopes should be marked for the attention of the Bulletin Group. *(Telephone: 071-601-4030)*

ISBN 1 85730 045 9
ISSN 0263-6123

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I: Introduction and Summary

This paper attempts to improve our understanding of the determinants of secondary market prices of LDC debt (for a sample of seven countries for which the market is reasonably liquid[1]). To do this we augment existing models with a term designed to capture economic influences on prices (the Bank of England's "matrix"[2] is used since this represents an economic evaluation of country performance). In doing so we also counter criticism that the matrix has no predictive powers,[3] for we find that it is useful in helping to explain movements in secondary market prices. Such a result suggests that the matrix might therefore be useful in predicting when debt repayment problems will arise, since the secondary market price is itself likely to be a good indicator of periods during which such problems arise. (For a demonstration of how prices can be used to anticipate future external financing difficulties, see Hajivassiliou (1989) - Appendix 1 provides a short summary.)

The next section of this paper considers a simple theoretical model of secondary market prices. Then, in section 3, we briefly discuss recent research, especially that carried out by Cohen and Portes (1990) since we wish to use their model as a benchmark. After pointing out a number of anomalies in the existing literature we address some of the problems with Cohen and Portes' results. Our own findings are presented in section 4. These are extended in section 5 to consider country-specific price movements. Finally section 6 contains our conclusions.

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1 Argentina, Brazil, Chile, Mexico, Poland, Venezuela and Yugoslavia.

2 The "matrix" is a set of credit scores used by the Bank in its supervisory role (relating to banks' lending to problem/highly-indebted LDCs).

3 Such criticism stems, firstly, from the proposition that the matrix results from what is essentially a "backward-looking" exercise and, secondly, from the fact that it is not founded on a (behavioural) model. For these reasons it has been suggested that the matrix scores are unlikely to have much, if any, predictive capability when it comes to forecasting debt repayment problems.

II: Theoretical Considerations

We start by considering a country which has borrowed an amount of debt D, and which has scheduled payments of P_1, P_2, ... (with payments being made at the end of each future period). Creditors considering the value of these payments will need to discount them to find their present value (V). We assume that the (risk-adjusted) interest rate, r, is used in this discounting exercise, so that;

$$V = \sum_{t=1}^{\infty} \frac{P_t}{(1+r)^t} \qquad (1)$$

If it is the case that the debtor is able to make payments sufficient to repay the debt one would expect the present value of these payments to equal the debt's face value.[1] In other words, the secondary market price (S) (which we define as the ratio of V to D) would, under these circumstances, be equal to one. If, however, it was thought that the debtors' ability to pay was insufficient for future payments to meet obligations then one would expect that V would be less than D. In these circumstances the secondary market price would be less than one. (Of course, for a secondary market to exist - within which creditors can buy or sell debt - then certain conditions will need to be met. For example, creditors might have different objectives which require them to adjust their portfolios.[2])

A number of factors are likely to be important in determining the payments that a debtor is able to make. Of prime importance is its ability to earn foreign currency. (Thus, in practice, most heavily indebted countries have had to run large trade surpluses in order to try and meet obligations.) In addition, however, payments are likely to be a function of (net) transfer payments, foreign direct investment, capital flows, the consumption and investment performance

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1 It has been argued that sometimes willingness to pay is an issue in addition to ability to pay (and, hence, that the present value of payments may not necessarily equal the face value of the debt even if the ability to pay was sufficient). An excellent discussion of the issues behind capacity versus willingness to pay is Cohen (1991). We return to the problem below.

2 For example, different creditors might have different views as to the payments they expect to receive or they might use different discount rates. In practice, debtors might also try to buy back debt if their views/discount rates are different from those of creditors. For them to do so, however, requires them to break the "pari passu" clause contained in most commercial lending arrangements.

of the economy and the country's balance sheet (ie both its asset and debt position are important).[1]

Before considering a model in which the precise means by which these factors affect repayments are outlined, we first wish to illustrate that, even with a very simple model, there is little reason to expect a unit elasticity of prices with respect to interest rates (a result upon which previous research has placed much importance). We then present a more general (and hence more realistic) model which is later used as a basis for our empirical work.

First we will assume, for simplicity, that debt service payments[2] are a fraction (p) of exports (X). Hence, we will re-write (1) as;

$$V = \sum_{t=1}^{\infty} \frac{p X_t}{(1+r)^t} \tag{2}$$

Making the assumption that exports grow at a constant rate (x) we are able to simplify (2) still further;

$$V = \frac{p X_1}{(1+r)} + \frac{p (1+x) X_1}{(1+r)^2} + \frac{p (1+x)^2 X_1}{(1+r)^3} + \ldots$$

$$= \frac{p X_1}{(r - x)} \tag{3}$$

In calculating (3) note that we are assuming $r > x$. (In the case that $r < x$ the debtor's payments are, in present value terms, infinite, so that any level of debt can be repaid in full within a finite time. Interestingly, under this assumption it is possible for a country to be solvent without it making any repayments - see Cohen (1985) or (1991) for more details.) For the case we consider, however, to find the secondary market price we can simply use (3) to substitute for V giving;

1 Kindleberger (1978) and Minsky (1982) show how balance sheet considerations can affect a country's debt repayment capability. For an empirical demonstration of how such factors affect rescheduling see Lloyd-Ellis, McKenzie and Thomas (1989).

2 Clearly we are using a very simple model. For example, we assume payments are known with certainty. (Cohen (1990) introduces uncertainty of payments in discussing secondary market discounts and notes that, when risk is allowed for, under certain circumstances the expected (discounted) payments can actually exceed the face value of the debt.) Furthermore, in more general models "x" could represent "payments capacity" - one might argue, for example, that the trade balance, rather than exports, is more relevant to capacity to pay.

$$S = \frac{p\,X_1}{D\,(\,r - x\,)} \tag{4}$$

Taking logarithms gives;

$$\ln(S) = \ln(p\,X_1 / D) - \ln(r - x) \tag{5}$$

Even if we were to make a simplifying restriction relating the growth rate of exports to the interest rate (say, that the former was equal to one-half the latter), it is clear that there is little reason to suspect a near unit elasticity of the secondary market price with respect to interest rates. (Obviously were we to assume a zero growth rate of exports we would generate the required result, but this is probably somewhat unlikely to occur in reality.) Taking "reasonable" values for r and x does, however, suggest that an interest rate rise would lead to a significant fall in the secondary market price.

Before considering the model we wish to use it is perhaps worth considering (5) in the light of the "stylised facts".[1] For example, take the case of the countries the World Bank define as "Severely Indebted Middle-Income Countries" (see World Bank (1990)). During 1980 their average long-term debt service to export ratio was close to 30%, their debt export ratio was close to 2, while interest rates averaged a little under 14%. For the secondary market price of these countries as a whole to have turned out less than one, then, according to (5), their average growth rate (of exports) would have had to been expected to be less than -1% (not a very likely scenario given that the average performance measured over the whole of the 1970s was a rise of close to 20% per annum). This rather suggests that, had there been a secondary market at this time, then very few creditors would have been willing to sell debt at less than its face value[2] (which perhaps helps explain why there was not a market!).

By the end of 1982, however, the situation was somewhat different - the average long-term debt service to export ratio had risen slightly, to 33%, the debt export ratio had risen to nearly 3, while the interest rate had remained close to 14%. Taken together these factors implied that the secondary market price would be less than one if the growth rate of exports was expected

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1 For a more detailed description of the evolution of LDCs debt problem over the past decade see Dicks (1991).

2 It is worth bearing in mind, however, that interest rates were very volatile during 1980. LIBOR reached close to 20% at one stage - implying that, had creditors expected such a high rate to continue, then the relevant growth rates (for which the secondary market price would have ended up less than one) would have been 5%.

to be less than 3%. In fact, the actual growth rate had averaged -1.3% p.a. during the period 1980 to 1982 - so that an expectation of zero growth in the near future would not have represented a marked break with the recent past at that time. (Such an outcome would have implied a secondary market price of close to 80%.) Thus it is not surprising that the secondary market began to develop in 1983 - creditors' views relating to debtors' future payment prospects are likely to have changed following the unusually low growth rate of exports between 1980-82 and the variance of expectations may have risen too (see also Stone (1990) and the references therein). By the time transactions reached significant levels (around 1986[1]), the debt export ratio had risen still higher (to 3.75 by the year-end), so that it was hardly surprising that rather more creditors then began to convert their claims (thus accepting that (at least some of) the interest and/or principal payments due them would not be received).

Since we have made a number of simplifying assumptions in deriving (5) we do not choose to use this model in our empirical work - it is likely to be too simple to be able to explain much more than the stylised facts. In addition to using interest rates we want to test whether or not a number of additional factors affect prices. In particular, we want to make use of the Bank's "matrix", which has been used for a number of years when supervising the UK banks.[2] For our purposes, the scores from this matrix are used to gauge the risks involved in lending to the countries involved - with a higher score indicating an increased risk (strictly speaking the matrix was designed to measure the ultimate recoverability of bank loans, but we will refer to this as "risk" for short). Originally the matrix identified fifteen indicators of debt repayment and servicing difficulties, assigning points to each according to their relative importance. In 1989 some minor changes were made to the matrix, with the total number of factors being increased to sixteen. Rather than make explicit the role perceived for each factor in affecting the repayment stream we begin by considering the aggregate scores as measures of risk. This we assume depends upon both payments capacity and willingness to pay.[3] Since we can think of a higher risk as involving either a reduced likelihood of creditors receiving a given stream

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1 The total volume of debt transactions in the secondary market rose from close to $2bn in 1984 to around $7bn in 1986 (World Bank Quarterly Review, 1989).

2 Basically we have used the matrix as a credit-scoring technique which takes into account three broad classes of factors related to rescheduling ("A" factors), arrears ("B" factors) and economic developments ("C" factors). Since secondary market prices is amongst the latter, then we remove this factor from the total scores when using the matrix to "explain" secondary market prices. Appendix 2 gives details of how the matrix is scored.

3 Such an assumption seems reasonable since the rescheduling and arrears factors can be thought of as measures of "reputation" (in addition to depending upon economic factors) whilst the economic factors are unlikely to depend upon willingness very much, if at all.

of payments or a reduction in the expected value of future payments then this suggests replacing (l) with;

$$E(V) = \sum_{t=1}^{\infty} \frac{E(P_t)}{(1+r)^t} \tag{6}$$

where E(.) denotes the expectations operator, and then letting the discounted stream of expected payments be a function of our matrix scores (MS);

$$E(V) = \alpha_0 MS^{\beta} \tag{7}$$

The matrix scores themselves will depend upon a large number of factors;

$$MS = f(Hr, a, r, D, X, M, Rs, Y, z) \tag{8}$$

where Hr is the country's history of rescheduling, a arrears, r interest rates, D debt stocks, X exports, M imports, Rs reserves, Y gdp, and z other factors. Note that f(.) is a non-linear function. Note also that, although the matrix is designed to capture interest rate effects (for example, through their influence on the interest service ratio), we might wish to augment (7) with an interest rate term (r) in order to facilitate a comparison with other models and to test if the weight assigned to interest rates in the matrix is "optimal" (in the sense of it being the best value to use in building a model to predict secondary market prices). Note too that interest rate volatility will be captured only to the extent that it influences the various matrix factors - no attempt is made to measure the variance of interest rates explicitly.

The use of matrix scores raises a few other issues. First, since the Bank's matrix scores are not made available to the commercial banks then, by using the former to explain secondary market prices, one is in effect testing to see whether or not official and market views of the likelihood and severity of debt problems coincide.[1] Second, there is a danger that higher matrix scores <u>cause</u> secondary market prices to fall simply because, by requiring higher provisions, this of itself leads to losses for the banks with regard to lending to particular debtors. Note, however, that this point is unlikely to be valid - both because the Bank does not impose its scores on banks - banks calculate their own scores, using them as a basis for discussion with supervisors - and because, anyway the UK banks comprise only a small proportion of total lending to LDCs (less than 15% for the main Latin American debtors).

1 Of course, even if they were made public one would still be testing to see whether or not official and market views coincide, but with a problem of direction of causation.

Moreover, UK banks have anyway often over-provided compared to our scores, so that they are rarely likely to need to alter provisions at precisely the same time as scores rise. Of course, even if one feels that there is a danger that matrix scores do "cause" prices this does not invalidate using them as an explanatory regressor.

We also augmented (7) with a dummy variable (CITI) designed to capture the effects of Citibank's announcement in May 1987 that it intended to make provisions (of close to $3bn) against its LDC exposure.[1] This we chose to do so as to encompass previous research (see, for example, the model estimated by Cohen and Portes (1990)). Although we recognise that it would be interesting to try to explain such a decision within the model, we felt it would be beyond the scope of this research to attempt to do so.[2] Adding CITI and interest rates to the model gives;

$$E(V) = \alpha_0 \, MS^{\beta 1} \, r^{\beta 2} \, CITI^{\beta 3} \qquad (9)$$

Dividing through by the face value of debt and taking logs gives an expression for the secondary market price;

$$\ln S = \ln \beta_0 + \beta_1 \ln MS + \beta_2 \ln r + \beta_3 \ln CITI \qquad (10)$$

where $\ln \beta_0 = (\ln \alpha_0 - \ln D)$. Obviously one needs to augment (10) with an error process for it to be estimable. Cohen and Portes, for example, add an error term which is assumed to be normally distributed. Before considering our own empirical efforts, based on (10), in the next section we briefly review recent research efforts.

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1 Chase Manhattan increased its loan loss reserves by $1.6 bn six days later and during June six other large US banks added $5.9 bn to LDC loan reserves, while UK banks began to provision during the middle of the month (see Stone (1990)).

2 Clearly one would need a model to analyse banks' behaviour, rather than one which identifies factors relevant to just one asset in their portfolios.

III: Previous Research

Although research using cross-section data has suggested that much of the variation in secondary market prices can be attributed to the underlying economic performance of the countries concerned (see, for example, Cooper (1990)), time-series models (such as that of Cohen and Portes (1990)) suggest that, for those countries for whom the secondary market can reasonably be characterised as "liquid",[1] prices of long-term debt are driven primarily by a set of "common factors" comprising interest rates (with a near unit elasticity) and a factor the authors term "systemic risk" (the latter not being correlated with macrovariables and therefore interpreted by the authors as "a set of factors that are common to the indebted countries only"). Such conflicting findings present something of a conundrum. How can cross-country variation in secondary market prices be determined by economic factors and yet the time-series variation in prices not be?

To some extent, the paradox has been addressed by Stone (1990) who found, using panel data for the period March 1986 to October 1989, that prices were sensitive to policy announcements associated with changes in future lending and to changes in key macroeconomic aggregates external to debtors (in line with the Cohen and Portes model), but not to innovations in LDC trade flows and reserves. Several problems in Stone's work are evident, however. First, the "external" macroeconomic factors that he found to be important and correctly signed did not include the interest rate - a somewhat surprising result, especially given the major role attributed to it by Cohen and Portes (and, of course, by the simple theoretical model outlined earlier). Second, the country-specific factors incorporated in Stone's analysis do not include many which Cooper identifies as important (for example, debt to GNP ratios and debt service payments to exports ratios[2]). Given these limitations it is perhaps not surprising to learn that the results of Stone's analysis leave a large proportion (generally more than 50%!) of the variance in secondary market returns unexplained.

Although not directly comparable, the results contained within Cohen and Portes hint that a much better fit of the data can be obtained if one first analyses "common trends" within different countries' prices and then seek to explain deviations from this trend (as regards country-specific data). Their work also suggests that it may be worth decomposing debt into

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1 Such a characterisation follows discussions with market traders (see Wilson (1989) for details). This paper also provides a useful survey of the secondary market's origins, its sources of supply and demand and the mechanisms through which trading take place.

2 Both of which affect our "matrix" scores, suggesting that, in the eyes of the supervisors at least, they could be relevant.

different maturities. (Such a decomposition can be justified on the basis that differing seniorities apply to different maturities.)

For these reasons we choose to follow the Cohen and Portes procedure of first estimating a model to explain the average price of long-term debt (using the same sample of countries that they studied). After replicating their work, the next section considers some of the problems evident with the model that they use and with their results. This leads us to make several small changes to the way the data are treated and to the functional form of the estimated model. It also leads us to emphasise the time-series properties of the data studied (leading us to use cointegration techniques).

IV: Results

The model we begin with is based on equation (10) from section II (which encompasses that estimated by Cohen and Portes (1990)). To begin with we suppress the role of the matrix, so as to consider a model identical to that of Cohen and Portes. They report a preferred model for the average price of long-term debt for the seven countries Argentina, Brazil, Chile, Mexico, Poland, Venezuela and Yugoslavia as;

$$\ln S_t = \underset{(-)}{5.99} - \underset{(0.17)}{0.96} \ln r_t - \underset{(0.06)}{0.20}\ CITI_t \qquad (11)$$

Sample period	1986.3 - 1989.11 (N = 45)
R^2	= 0.80
σ	= 0.03
DW	= 0.49
Linearity $X^2(1)$	= 6.14
Homoscedasticity $X^2(1)$	= 0.53
Normality $X^2(2)$	= 9.23

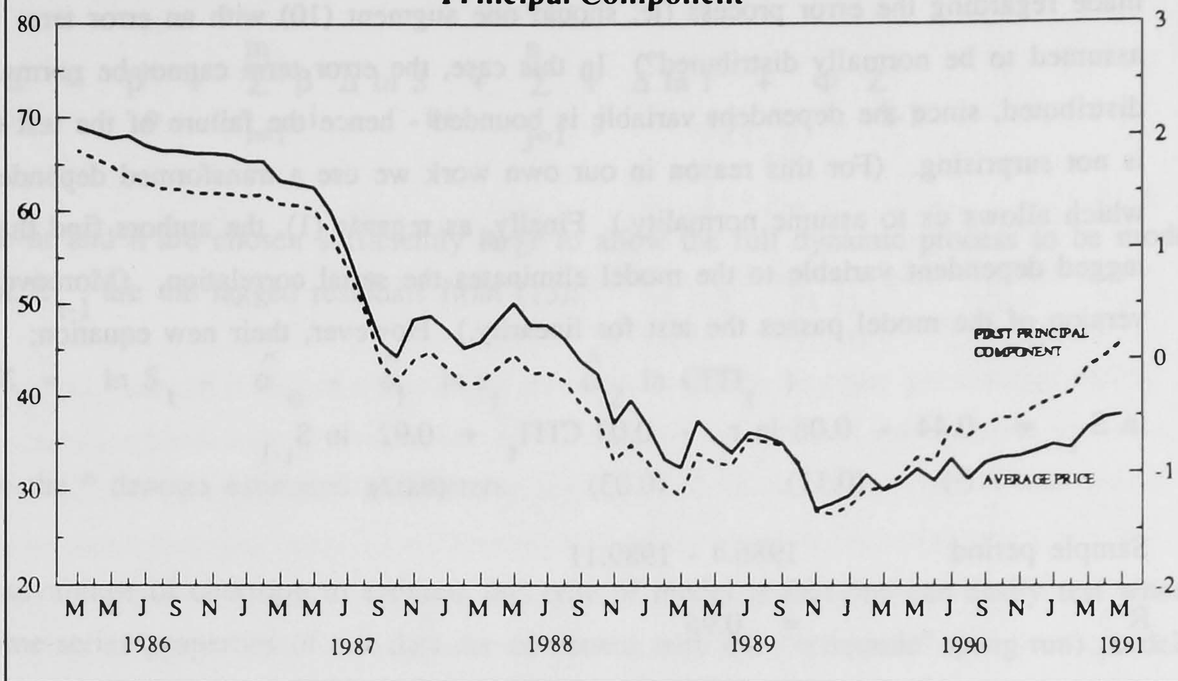
where S is the average price (weighted by debt stocks[1]), r is LIBOR and CITI is a dummy taking the value 1 from May 1987 onwards.[2] Principal components had been used to justify the decision to model the average price - it turned out that 86% of the variance in prices was explained by the first principal component and that this component was, as Cohen and Portes report, "an almost perfect average of all prices". Chart 1 shows the two series (from our replication of their work), serving to highlight their high correlation.

1 The weights used are based on end-year data relating to total debt stocks (ie including both official and commercial debt). The source for these data is the World Bank Debt Tables (1988-89).

2 The monthly price data are from Salomon Brothers. Standard errors are shown in brackets. Note that there is little point in taking logs of CITI since this would merely result in switching ones for zeros and vice versa.

Chart 1

Average Prices and the First Principal Component



Before considering our own results a few points are worth making relating to the Cohen and Portes model. First, consider the near unit elasticity of prices with respect to interest rates - a finding which Cohen and Portes report as a "very robust result". Although this is true, in the sense that the estimated coefficient is very near to one, in fact an F-test to test the restriction is actually rejected (by a substantial margin at the 95% significance level). Besides, as regards theoretical reasoning, it is clear that a strong assumption is needed in order to generate the unit elasticity result - each and every country has to be in the position that it is servicing its debt independently of its face value (see section V of Cohen and Portes). In our simple model such a result holds only when the growth rate of exports is deemed equal to zero (see section II).

Second, one needs to consider the model's test results a little more closely. In particular, one notices three problems with the Cohen and Portes model;

(1) the low DW statistic indicates significant (first-order) serial correlation,

(2) the failure of the test regarding linearity,

(3) the failure of the test regarding normality of the residuals.

Taking the problems in reverse order (3) may not be very important - it could simply reflect the two very large residuals which occurred in May 1987 and November 1989. Taken together with (2), however, one is naturally led to question the assumptions that Cohen and Portes have made regarding the error process (ie. should one augment (10) with an error term which is assumed to be normally distributed?) In this case, the error term cannot be normally distributed, since the dependent variable is bounded - hence the failure of the test for linearity is not surprising. (For this reason in our own work we use a transformed dependent variable which allows us to assume normality.) Finally, as regards (1), the authors find that adding a lagged dependent variable to the model eliminates the serial correlation. (Moreover, this version of the model passes the test for linearity.) However, their new equation;

$$\ln S_t = \underset{(-)}{0.44} - \underset{(0.11)}{0.06} \ln r_t - \underset{(0.03)}{0.03}\, CITI_t + \underset{(0.07)}{0.92} \ln S_{t-1} \qquad (12)$$

Sample period	1986.4 - 1989.11
R^2	= 0.96
σ	= 0.014
Durbin's λ	= 1.42
Linearity $X^2(1)$	= 0.04
Homoscedasticity $X^2(1)$	= 9.80
Normality $X^2(2)$	= 7.03

clearly suffers from a number of new problems. One is the failure of the test relating to homoscedasticity. The second (and much more important) is the insignificance of both the interest rate term and the CITI dummy variable. Once this is recognised, it becomes apparent that (12) is really a simple time-series model which predicts that prices in any month will be 92% of what they were the previous month (plus a constant). It provides no justification for the role of any "economic" variable.

One obvious way forward in trying to improve upon the two Cohen and Portes' models would be to estimate a generalised error-correction model of secondary market prices (see Banerjee, Galbraith and Dolado (1988)). An alternative procedure (which we later choose to follow) would be to estimate a two-step Granger-Engle model in which the first (levels) equation describes the long-run characteristics of the model;

$$\ln S_t = \alpha_o + \alpha_1 \ln r_t + \alpha_2 \ln CITI_t \qquad (13)$$

and the second (dynamic) equation explains how prices evolve over time;

$$\Delta \ln S_t = \beta_o + \sum_{i=1}^{m} \beta_i \Delta \ln S_{t-i} + \sum_{j=1}^{n} \Psi_j \Delta \ln r_{t-j} + \Phi Z_{t-1} \qquad (14)$$

where m and n are chosen sufficiently large to allow the full dynamic process to be modelled whilst Z_{t-1} are the lagged residuals from (13);

(ie $Z_t = \ln S_t - \hat{\alpha}_o - \hat{\alpha}_1 \ln r_t - \hat{\alpha}_2 \ln CITI_t$)

where the ^ denotes estimated parameters.

One advantage of choosing to estimate this type of model is that one can easily test whether the time-series properties of our data are consistent with our "economic" (long-run) model. This amounts to testing whether the variables in (13) above form a cointegrating vector (see Hendry (1986) for an introduction to cointegration theory). Thus, before estimating our own model of secondary market prices we first consider whether or not the Cohen and Portes "levels" equation passes the tests for cointegration.

Having first considered the time-series properties of the variables we are working with, (all of which appear to be integrated of order one (see Appendix 3 for details), we next replicated the Cohen and Portes long-run equation, discovering that prices, the interest rate and the CITI dummy do not form a cointegrating vector (since the Dickey-Fuller (DF), Augmented Dickey-Fuller (ADF) and the Cointegrating Regression Durbin Watson (CRDW) test results are all failed (being too close to zero)). Our results based on an extended sample period (taking into account more recent data) were as follows;[1]

$$\underset{}{\ln S} = \underset{(0.374)}{5.196} - \underset{(0.198)}{0.533} \ln r - \underset{(0.068)}{0.422} CITI \qquad (15)$$

1 Note that similar test results were obtained when using the same sample period as Cohen and Portes (with the DF test equal to -1.45 and the ADF -1.00). We use the longer sample period throughout our work since it gives us more degrees of freedom.

Sample period	1986.3 - 1991.5 (N = 63)
R^2	= 0.66
σ	= 0.17
CRDW	= 0.23
DF	= -1.85
ADF	= -1.14

where we have dropped the time subscripts for convenience.

Before considering ways in which (15) can be improved it is worth emphasising the overall standard error of this model. At close to 17% (11% for the shorter sample period used by Cohen and Portes) we obtain a worse goodness-of-fit than that suggested by Cohen and Portes. Since in every other respect - in terms of estimated coefficients, standard errors of the coefficients, R^2 of the equation and test statisitics - we were able to obtain very similar results to those of Cohen and Portes, it is important to note that our estimate (of 17%) is the correct bench-mark against which other models need to be gauged (when searching for a better explanation of secondary market prices).[1]

Before adding the matrix scores to the model we first changed the way prices were weighted together to take into account two problems evident with the Cohen and Portes dataset - first, that the weights they used covered all debt, when we felt that it was only commercial debt which is appropriate (since only this is traded)[2] and, second, that the weights they used were not interpolated (and so "jumped" between each December and January). Table 1 illustrates the importance of the latter point, showing the amounts of debt outstanding at the end of 1989 by creditor. In practice, however, we found that neither of these changes made any significant difference to our results (although the estimated coefficients did change slightly) - the model used by Cohen and Portes does not appear to be a cointegrating vector (and so should be rejected).

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1 We would like to thank Daniel Cohen for verifying the discrepancy between our results.

2 This is particularly important in the case of Poland which, through much of the sample period, had a large proportion of its debt owed to official creditors. As a result the Cohen and Portes average price series gives a much higher (relative) weight to Polish debt than ours does (or compared, for example, to the average price series published by Salomons).

Table 1: 1989 Long Term Debt Outstanding and Disbursed by Creditor ($ mn)

Creditor	Country						
	Arg.	Bra.	Chi.	Mex.	Pol.	Ven.	Yug.
Official	9,896 (18.6)	24,345 (27.0)	4,604 (32.9)	16,786 (20.9)	23,811 (68.5)	959 (3.2)	6,678 (37.6)
of which							
Multilateral	4,351 (8.2)	11,088 (12.3)	3,550 (25.4)	10,753 (13.4)	496 (1.4)	560 (1.9)	2,660 (15.0)
Bilateral	5,546 (10.4)	13,257 (14.7)	1,054 (7.5)	6,033 (7.5)	23,315 (67.1)	399 (1.3)	4,018 (22.6)
Commercial Banks	29,122 (54.7)	53,251 (59.0)	8,797 (62.8)	53,989 (67.3)	9,268 (26.7)	25,269 (84.6)	10,977 (61.7)
Other	14,211 (26.7)	12,696 (14.1)	596 (4.3)	9,481 (11.8)	1,668 (4.8)	3,634 (12.2)	129 (0.7)
Total	53,229 (100.0)	90,292 (100.0)	13,997 (100.0)	80,256 (100.0)	34,747 (100.0)	29,862 (100.0)	17,784 (100.0)

Notes: (1) Figures in brackets are percentages of the total.
(2) Source: World Debt Tables 1990-91.

Next we tried augmenting (15) with the (total) matrix scores, denoted MS, (taken from the Bank matrix and weighted together using commercial banks' debt stocks), finding that the new variable (MS not ln MS)[1] was not only correctly signed and significant but that the resulting vector passed the tests for cointegration (note that the ADF test result is irrelevant since the correlogram indicates that no lagged values of changes in residuals were significant in the ADF (residuals) regression);

$$\ln S = \underset{(0.210)}{6.363} - \underset{(0.105)}{0.918} \ln r - \underset{(0.040)}{0.149} CITI - \underset{(0.001)}{0.014} MS \qquad (16)$$

Sample period	1986.3 - 1991.5
R^2	= 0.91
σ	= 0.09
CRDW	= 0.65
DF	= -3.77
ADF	= -3.98

To counter the problem associated with the fact that it is inappropriate to assume normality we next considered the same model but with the dependent variable first transformed, using the logit transformation;

$$\ln TS = \ln \frac{(S / 100)}{1 - (S / 100)} \qquad (17)$$

(Note that since our prices are in cents per dollar they are first scaled so as to be between zero and one.) The regression results based on the logit model again provide strong support for the notion that the matrix contains useful information for helping to explain movements in secondary market prices;

$$\ln TS = \underset{(0.403)}{4.365} - \underset{(0.201)}{1.563} \ln r - \underset{(0.077)}{0.456} CITI - \underset{(0.002)}{0.024} MS \qquad (18)$$

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1 This slight departure from (9) (ie. the decision to use the actual matrix scores rather than their logged values) was made because, later in the paper, we want to split the total scores into their components and such a procedure facilitates comparison between results (allowing, for example, tests of linear restrictions). In practice, the decision to refrain from using logged values is not important - we found that when MS was replaced with ln MS the new term was still both highly significant and correctly signed, and the tests for cointegration still passed.

Sample period 1986.3 - 1991.5

R^2 = 0.92

σ = 0.17

CRDW = 0.64

DF = -3.77

ADF = -4.09

More importantly, it is clear from the DF and ADF test results that (18) is a valid representation of a long-run equation (ie. it is a cointegrating vector). Comparing (16) and (18) note the bigger (absolute) coefficient on interest rates in the latter. Although the former is an elasticity, for (18) the elasticity is given by the expression -(1-(S/100))*1.563. Calculated at the mean value of S for our sample this gives a figure of -0.85, only slightly smaller in absolute terms than for (16).

In order to investigate which of the three main classes of factors were relevant (see footnote 2 on page 5) we next tried splitting MS into three components - MSA representing the "A" factors, MSB the "B" factors and MSC the (economic) "C" factors. All three were then incorporated in the model when searching for a cointegrating vector. The "B" factors were found to be insignificant and so were dropped from the model, giving as our preferred long-run equation;

$$\ln TS = \underset{(0.371)}{4.121} - \underset{(0.198)}{1.208} \ln r - \underset{(0.072)}{0.434} \, CITI - \underset{(0.009)}{0.039} \, MSA - \underset{(0.015)}{0.057} \, MSC \qquad (19)$$

Sample period 1986.3 - 1991.5

R^2 = 0.93

σ = 0.15

CRDW = 0.67

DF = -3.83

ADF = -4.16

Note that, when the matrix score is disaggregated into its factors, the elasticity of the price with respect to LIBOR decreases in absolute terms to -0.66, whilst the coefficients on MSA and MSC are larger (more negative) than that on MS. Since the latter may well be capturing interest rate effects to some extent (the "C" factors include the interest service ratio, for example), it is unclear whether or not the overall effect of interest rates on prices is significantly different between (19) and (18).

These results show that the "A" and "C" factors (related to rescheduling and economic considerations respectively) are each important determinants of secondary market prices (more especially the latter). That "B" factors do not matter is perhaps not too surprising since it is likely that the process of deterioration of a country's finances is likely to involve both "B" and "C" factors rising simultaneously (or, at least, almost so).[1] It may also reflect the fact that information relating to arrears is less easily available to those trading in the secondary markets than are economic data or details of reschedulings.

As regards the role of "A" factors, it is worth noting that they attract a negative coefficient, since there are circumstances under which a rescheduling would raise the value of existing debt (if, for example, it involved a write-down or forgiveness of existing debt - as occurred recently in the case of Poland's Paris Club deal). One might argue that this is particularly likely if creditors had already witnessed the debtor facing payment difficulties (perhaps already captured in the "C" (economic) factors), for then, when a more generous (or earlier) than anticipated rescheduling occurs, the value of existing claims (in the eyes of creditors) may actually be raised somewhat - particularly if the rescheduling results in a reduction in the "debt overhang" (see Krugman (1989) or Sachs (1989) for an exposition of the debt overhang argument). Moreover, when the rescheduling is a Paris Club affair (ie official debt is being rescheduled/written down) it is possible for there to be no overall improvement in the debtor's outlook but for there to still be an increase in the value of private sector claims (in effect, by rescheduling, the official sector would be extending insurance to the banks' bad debt (or equivalently through their making new money available) to badly performing debtors).

Of course, there are also circumstances under which the opposite result holds (if, for example, reschedulings or (perhaps more realistically) declarations of moratoria arrive "out-of-the-blue", thus representing unexpected "bad" news for creditors).[2] The conclusion one must draw on the basis of our equation is that the majority of those trading in the secondary market must take a

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1 It is also worth pointing out that "B" factors were found to be significant when it came to explaining prices for individual countries (Venezuela, for example). Section V below gives details.

2 Take, for example, the sharp decline in the secondary market prices of Brazil and Ecuador following their unexpected announcement of interest payment suspensions in the first quarter of 1987.

pessimistic view of reschedulings (and/or that they are consistently less generous than expected) - writing down the value of existing claims as a result.

Having produced a reasonable long-run model we next proceeded to estimate a full dynamic model of secondary market prices (on the basis of (14) but where this equation is augmented with difference terms in each of the three matrix scores (MSA, MSB and MSC)). Having tested down from a general model (in which lags of up to four months were permitted) our preferred model took the form;

$$\Delta \ln TS_t = \underset{(0.010)}{-0.012} - \underset{(0.240)}{0.516}\,\Delta \ln r_t + \underset{(0.244)}{0.529}\,\Delta \ln r_{t-1} - \underset{(0.043)}{0.159}\,\Delta MSC_t \quad (20)$$

$$+ \underset{(0.041)}{0.087}\,\Delta MSC_{t-1} - \underset{(0.027)}{0.056}\,\Delta MSC_{t-3} - \underset{(0.082)}{0.449}\,Z_{t-1}$$

Sample period		1986.8 - 1991.5
R^2	=	0.41
σ	=	0.075
DW	=	1.95
Autocorrelation $F(4,47)$	=	0.80
Normality $X^2(1)$	=	0.75
Linearity $X^2(4)$	=	4.13
Heteroscedasticity $X^2(1)$	=	1.71
Parameter stability[1] $X^2(6)$	=	3.94
Parameter stability[1] $X^2(9)$	=	6.72
Parameter stability[1] $X^2(12)$	=	11.32

This model not only explains the past fairly well, as is illustrated by charts 2, but passes the forecast tests (ie. the parameter stability tests). Although the residuals pass all the diagnostic tests used, several large residuals occurred at times when we know that other factors than those incorporated in the matrix were relevant. Thus we found that by including three additional

1 Note that these tests are based on models estimated using shorter sample periods.

dummy variables as regressors (to "explain"[1] three large changes in price which occurred but were not picked up by the model) we could reduce the standard error of the model from 7 1/2 % to a little over 5 % (though at the cost of reducing the role of changes in interest rates in explaining price changes). The first such variable we found to be significant took the value 1 in 1989.2 and -1 in 1989.4 representing speculation around the time the Brady Plan was announced. Prices for Brazilian debt fell particularly fast in February (when the Brazilian President announced that interest payments due in March might be missed), though a number of other countries prices fell sharply too (particularly after the US Treasury Secretary, Nicholas Brady, postponed until March his planned announcement of new proposals to deal with the debt crisis). Despite mixed reviews by April it was clear that, in the words of one senior banker, the "Brady Plan" would be "net positive for creditor banks" (see IFR April 15).

The second dummy used took the value 1 in November 1988 and November 1989 - the month during which US banks tend to make year-end adjustments to their portfolios. In November 1988 the market suffered what was described by some commentators as a "free fall" (see IFR November 12) as regional banks tried to clear their books of mainly Latin American debt. Prices for both Brazilian and Mexican debt fell by close to 10% in one week (this despite a $5.2bn new money package for the former). Towards the end of the month rumours that Brazil would suspend its debt/equity auctions altogether (in order to concentrate on cutting inflation) added to worries. November 1989 also witnessed large-scale selling by banks for end-year bookkeeping purposes, leading to prices falling sharply (see IFR November 11).

The third dummy found to be significant took the value 1 from July 1990 onwards, this being the date at which our Mexican price series jumps by more than ten percentage points. This might be a consequence of the Brady deal which reduced Mexico's bank debt significantly. It might, however, simply be the consequence of our having had to switch sources for our data at this point (see Section V below for a more detailed discussion).

In addition to showing how secondary market prices can be influenced by variables which theory suggests could be important determinants, (20) clearly fits the data very well - the standard error of the equation is less than one half that of the Cohen and Portes model (see

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1 Of course, the dummies do not, in a sense, "explain" anything. Hence the fewer the better (see Spanos (1986) for details of why one should avoid being too liberal in the use of dummy variables). However, as he makes clear, they are necessary if the model is to pass the tests for normality - we therefore used the minimum number of dummies necessary to pass these tests.

Chart 2

Actuals and Fitted from the Dynamic Model

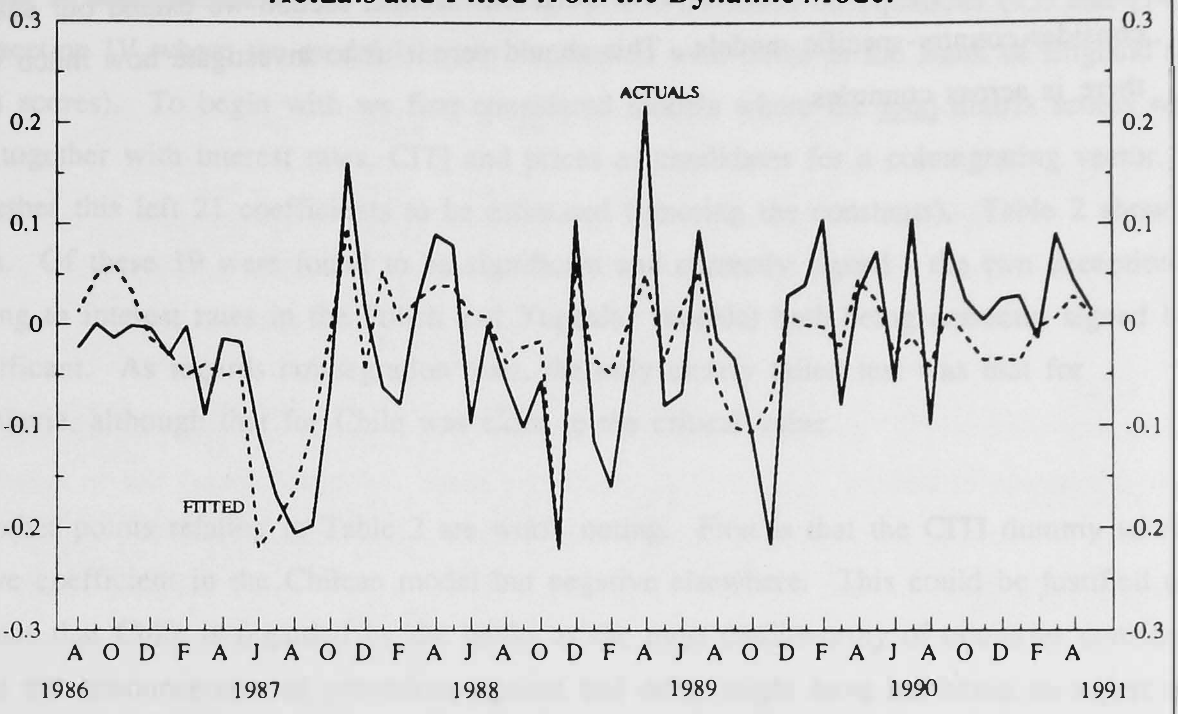
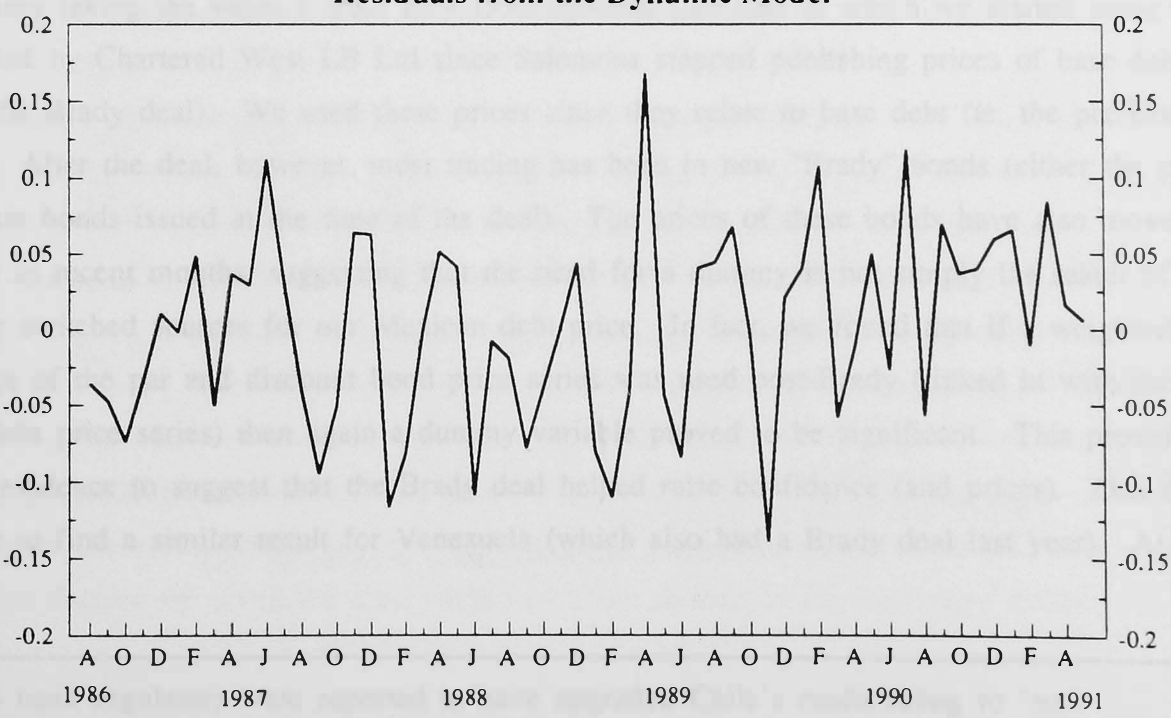


Chart 2

Residuals from the Dynamic Model



(15) above) and close to two thirds that of a simple time-series model (in which ln TS is explained solely in terms of lags of itself). Thus, the main conclusion of our work thus far is that, not only is it possible to find a role for "economic" terms in explaining secondary market prices, but they appear to do so fairly well. In the next section we extend our analysis to consider country-specific models. This should permit us to investigate how much variation there is across countries.

V: Country-Specific Models

In this section we report the results obtained from estimating two-step models of secondary market prices (again using the Granger-Engle estimation technique) keeping to the same model structure as was used earlier to explain average prices (ie based on equations (13) and (14) from section IV where the model is again augmented with terms in the Bank of England's matrix scores). To begin with we first considered models where the total matrix scores were used, together with interest rates, CITI and prices as candidates for a cointegrating vector. Altogether this left 21 coefficients to be estimated (ignoring the constants). Table 2 shows our results. Of these 19 were found to be significant and correctly signed - the two exceptions (relating to interest rates in the Polish and Yugoslav models) both being correctly signed but insignificant. As regards cointegration tests, the only clearly failed test was that for Yugoslavia, although that for Chile was close to the critical value.

Two other points relating to Table 2 are worth noting. First is that the CITI dummy takes a positive coefficient in the Chilean model but negative elsewhere. This could be justified on the basis that Chile is regarded by the banks as the most creditworthy of countries considered, so that the announcement of provisions against bad debts might have led banks to adjust their portfolios away from Brazil, Mexico etc.. towards Chile.[1] Alternatively it could indicate mis-specification.[2] Second, is that we found it necessary to augment the Mexican model with a dummy taking the value 1 from July 1990 onwards (the date at which we started using data provided by Chartered West LB Ltd since Salomons stopped publishing prices of base debt after the Brady deal). We used these prices since they relate to base debt (ie. the pre-Brady debt). After the deal, however, most trading has been in new "Brady" bonds (either the par or discount bonds issued at the time of the deal). The prices of these bonds have also moved higher in recent months, suggesting that the need for a dummy is not simply the result of our having switched sources for our Mexican debt price. In fact, we found that if a weighted average of the par and discount bond price series was used post-Brady (linked in with the old base-debt price series) then again a dummy variable proved to be significant. This provides some evidence to suggest that the Brady deal helped raise confidence (and prices). One might expect to find a similar result for Venezuela (which also had a Brady deal last year). At first

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1 US bank regulators were reported to have upgraded Chile's credit rating to "non-restructuring status" in 1990. This year Salomons suggested that Chile had become Latin America's first "investment-grade credit" (see Salomon Brothers (1991)).

2 If, for example, the matrix were to have a bias towards only recording "negative" factors, this might explain why the model for Chile has the positively signed CITI dummy.

Table 2: Individual Country Results: Long-Run Equations with Total Matrix Score

Dependent Variable: ln $TS_{i,t}$ (where i refers to country and t to time).

Sample Period: 1986.3 - 1991.5 (N = 63)

Explanatory Variables	Arg.	Bra.	Chi.	Mex.	Pol.	Ven.	Yug.
Constant	5.585 (9.01)	6.283 (7.90)	4.453 (10.78)	3.304 (9.68)	2.297 (2.68)	6.102 (10.71)	2.740 (3.24)
ln r_t	-2.094 (6.53)	-2.258 (5.76)	-1.233 (5.31)	-1.244 (6.91)	-	-2.377 (7.62)	-
$CITI_t$	-0.920 (7.52)	-0.699 (4.76)	0.344 (4.58)	-0.107 (1.70)	-0.540 (3.72)	-0.478 (4.55)	-0.762 (3.59)
$MATOT_t$	-2.372 (11.30)	-2.930 (9.56)	-6.013 (8.21)	-2.413 (5.97)	-2.967 (8.61)	-1.655 (3.75)	-2.897 (3.28)
NMEXDUM	-	-	-	0.618 (10.96)	-	-	-
Statistics							
R^2	0.92	0.87	0.71	0.86	0.65	0.83	0.74
σ	0.28	0.30	0.19	0.14	0.36	0.26	0.33
DF	-3.68	-2.99	-1.15	-3.80	-2.04	-3.50	-1.16
ADF	-3.85^*	-3.33^*	-2.87^*	-4.48^*	-3.41^*	-3.95^*	-1.66^*
CRDW	1.75	1.63	1.17	1.50	1.04	1.46	1.80
Interest rate elasticity	-1.417	-1.257	-0.428	-0.607	-	-0.105	-

Notes: (1) Figures in brackets are t-statistics. As regards the DF and ADF tests, a * denotes which test is the more relevant.
(2) Matrix score coefficients are rescaled by a factor of 100.

blush prices appeared to have risen post-Brady. When a dummy was added to the long-run model, however, we found it to be insignificant.

Given these (fairly encouraging) results we next tried splitting the matrix scores into their three components. Table 3 shows our preferred models, obtained after testing down from an equation which permitted roles for all three groups of factors. As with the average-price model we found that, with the exception of Yugoslavia, by adding matrix scores we were always able to improve the models - both in the sense of the new term(s) being significant but also (and more importantly) in terms of improving the test results for cointegration. Of the countries considered all but the Yugoslavian model passed the tests and so represent satisfactory long-run equations.

As regards the explanatory variables which turned out to be relevant, "C" factors were the most useful of the matrix variables, being significant in 6 of the 7 models (Venezuela being the exception) and always attracting a negative coefficient. In all of the models we were also able to find a role for either "A" or "B" factors (and both in the case of Brazil).[1] As regards the estimated coefficients those relating to the "C" factors are usually close to 6, though with Mexico being significantly smaller (though, interestingly, the discrepancy is reduced if the Brady dummy is excluded). "C" factors also have a smaller effect in the Polish equation (where "A" factors play a much larger role than in the other models) and Venezuela (where we found the estimated coefficient, at -0.2, insignificantly different from zero).

The estimated coefficients pertaining to "A" factors vary significantly, with Polish prices particularly sensitive to changes in these scores. This is perhaps not too surprising given our comments earlier regarding the nature of reschedulings/refinancings - it is sometimes hard to judge whether or not such an event is good or bad for the debtor's repayment prospects let alone gauge how important. "Voluntary" refinancing, for example, is likely to be viewed by creditors somewhat differently from a (more typical) "involuntary" rescheduling. Similar arguments might be used to argue that the coefficients related to "B" factors could also vary somewhat. In fact, however, we find that, for those countries where we can hope to estimate coefficients the estimated effects are generally fairly small - between 1 and 2.5. This suggests that these factors are given the least weight by those dealing in the secondary markets.

1 Moreover, it should be noted that for two of the countries examined the "B" factor scores did not vary within the sample period considered. Hence they could not be included as separate regressors in these models (their effect being picked up in the constant).

Table 3: Individual Country Results: Long-Run Equations with Disaggregated Matrix Score

Dependent Variable: ln $TS_{i,t}$ (where i refers to country and t to time).

Sample Period: 1986.3 - 1991.5 (N = 63)

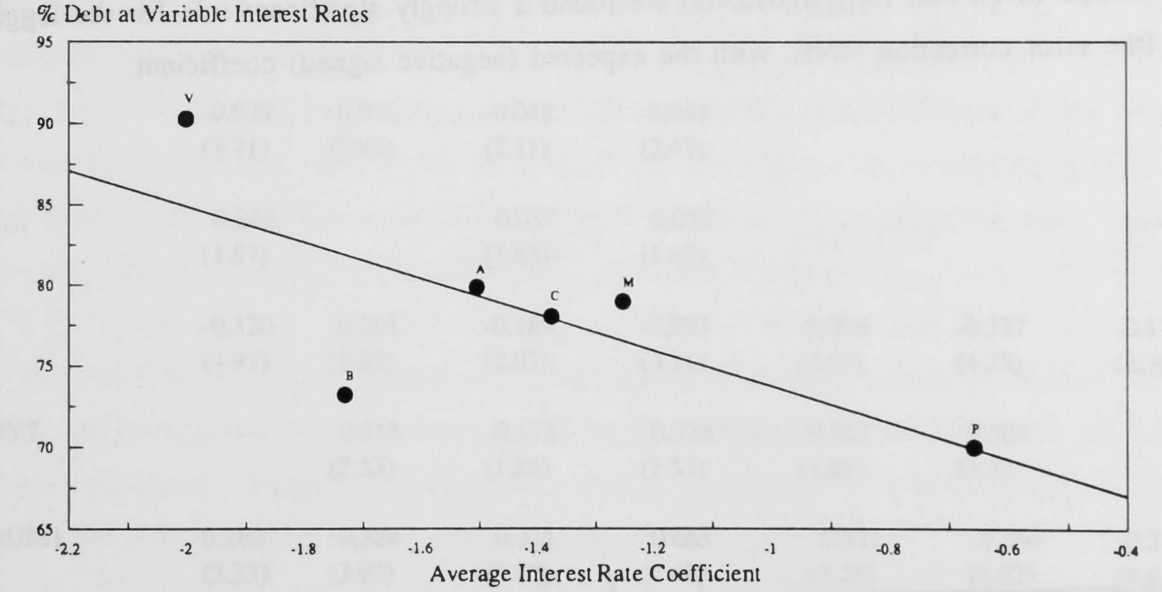
Explanatory Variables	Country						
	Arg.	Bra.	Chi.	Mex.	Pol.	Ven.	Yug.
Constant	4.446 (6.41)	5.715 (7.54)	4.595 (10.96)	3.305 (9.63)	3.522 (5.31)	5.099 (9.27)	2.170 (9.74)
ln r_t	-1.505 (4.12)	-1.731 (4.42)	-1.376 (5.54)	-1.254 (6.89)	-0.657 (2.06)	-2.001 (6.71)	-
$CITI_t$	-0.893 (7.43)	-0.645 (3.64)	0.281 (3.29)	-0.112 (1.74)	-0.326 (2.38)	-0.646 (6.80)	-0.826 (5.35)
$MSA_{i,t}$	-	-5.334 (2.92)	-3.444 (1.87)	-2.702 (4.06)	-9.656 (7.46)	-	-4.947 (2.96)
$MSB_{i,t}$	-1.814 (1.99)	-1.174 (1.79)	-	-	-	-2.409 (5.54)	-
$MSC_{i,t}$	-5.801 (4.81)	-5.969 (3.38)	-7.020 (7.14)	-1.858 (1.70)	-4.008 (2.67)	-	-5.987 (3.81)
BRADYMEX	-	-	-	0.647 (8.35)	-	-	-
Statistics							
$\bar{R}^2$	0.92	0.89	0.71	0.86	0.80	0.86	0.76
σ	0.27	0.28	0.19	0.14	0.27	0.23	0.31
DF	-3.76	-3.04	-1.42	-3.84	-2.34	-3.81	-1.33
ADF	-4.23*	-3.40*	-3.19*	-4.47*	-3.27*	-4.32*	-1.37*
CRDW	0.73	0.50	0.37	0.68	0.28	0.61	0.23
Interest rate elasticity	-1.019	-0.964	-0.477	-0.612	-0.430	-0.884	-

Notes: (1) Figures in brackets are t-statistics. As regards the DF and ADF tests, a * denotes which test is the more relevant.
(2) Matrix score coefficients are rescaled by a factor of 100. BRADYMEX is a dummy in the Mexican model only.

As regards interest rate effects, the estimated coefficients varied between 1 1/4 and 2 for the Latin American economies, but were much smaller for Poland and Yugoslavia (at close to 1/2, with the latter proving to be insignificant). One explanation for the divergence could be differing degrees of capital controls. For example, following an interest rate rise in developed countries an "open" economy (ie one without capital controls), facing an increase in its debt-service requirements, would tend to raise domestic interest rates, thus making the necessary adjustment to its trade balance. In contrast to this, a "closed" economy (where capital is imperfectly mobile) might not follow interest rates up and, hence, would not improve its trade performance. A second reason for the divergence may be that the proportion of debt which is variable rate or concessional will vary across countries. Chart 3 shows that the former explains some of the variation in the estimated interest rate effects for the countries we examined (the proportion of debt which is variable rate turned out to be a significant factor (t-value of 3.9) in explaining differences in the estimated interest rate coefficients).

Chart 3

Percentage of Total Debt at Variable Interest Rates Against Average Interest Rate Coefficient in Disaggregated Matrix Model



Around two-thirds of the variation in the estimated coefficients across countries was explained by this regression. Another factor which might be relevant in explaining the variation is the proportion of debtor's export markets accounted for by the US or by countries which keep their interest rates in line with the US.[1] Certainly the two economies where we found interest effects were small tend not to trade very much with the US (suggesting that it might be worth considering additional (non-US) interest rate terms in the models for these countries' prices).

As regards the Chilean long-run model we again find that the sign on CITI is positive (and therefore perhaps difficult to justify - though see above for one attempt). We found, however, that inclusion of the dummy was not significant to our cointegration test results - even without CITI the ADF test was passed. Thus the main conclusion of our country-specific long-run equations backs up our previous findings - adding matrix scores as regressors allows one to improve markedly on a model which uses only interest rates to explain secondary market prices.

Next we estimated some dynamic country-specific models using the lagged residuals from the levels regressions as error-correction terms.[2] Table 4 shows our results (obtained after testing down from the general model with at most 2 lagged values for each of the regressors[3]). Generally the models fit the data fairly well and exhibit desirable properties (for example, they generally appear to forecast well - see the tests for parameter stability). For each of the models (even that for Yugoslavia) we found a strongly significant role for the lagged residuals (the error correction term), with the expected (negative signed) coefficient.

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1 The argument here would be that if one country has the US as a major export market then when US interest rates rise then not only does the debtor face higher debt-service payments but its export market may also contract.

2 Of course, this is not appropriate in the case of Yugoslavia - since we have not found a cointegrating vector for this country's price. We should not, therefore, be surprised if the dynamic model for Yugoslavia is less stable than those of the Latin American economies' prices.

3 Had we started with general models with longer lags we might conceivably have improved the models fit slightly, but we felt that by doing so we might be in danger of over-fitting the data.

Table 4: Individual Country Results: Dynamic Equations

Dependent Variable: $\Delta \ln TS_{i,t}$ (where i refers to country and t to time).

Sample Period: 1986.7 - 1991.5 (N = 59)

Explanatory Variables	Country: Arg.	Bra.	Chi.	Mex.	Pol.	Ven.	Yug.
Constant	-0.044	-0.045	-0.005	-0.004	-0.017	-0.029	-0.013
	(2.90)	(2.75)	(0.04)	(0.37)	(1.78)	(2.51)	(1.27)
$\Delta \ln TS_{i,t-1}$	-	-	0.528	-	0.379	0.226	0.173
			(3.94)	-	(5.20)	(2.36)	(2.07)
$\Delta \ln r_t$	-0.527	-0.599	-	-	-	-0.582	
	(1.48)	(1.67)				(2.21)	
$\Delta MSA_{i,t}$	-	-	-	-	-0.028	0.025	-
					(3.01)	(3.18)	
$\Delta MSA_{i,t-1}$	-	-	-	-0.044	-	-	-
				(2.57)			
$\Delta MSB_{i,t-1}$	-				0.007	-	-
					(2.15)		
$\Delta MSC_{i,t}$	-0.029	-0.070	-0.048	-0.044			
	(1.71)	(2.89)	(2.11)	(2.47)			
$\Delta MSC_{i,t-1}$	0.028	-	0.037	0.039			-
	(1.87)		(1.63)	(1.88)			
$Z_{i,t-1}$	-0.320	-0.255	-0.189	-0.293	-0.096	-0.237	-0.116
	(4.91)	(3.83)	(2.07)	(3.11)	(2.61)	(4.25)	(3.36)
BRADYT	-	0.313	0.173	0.228	0.262	0.304	-
		(2.53)	(1.82)	(2.57)	(3.69)	(3.39)	
OTHDUM1	0.262	0.324	0.311	0.665	0.521	0.366	-0.370
	(2.33)	(3.63)	(3.29)	(7.48)	(7.26)	(4.02)	(4.65)
OTHDUM2	0.326	-0.315				-	-0.525
	(2.87)	(2.57)					(6.59)

Table 4 cont.

Statistics

	Country						
	Arg.	Bra.	Chi.	Mex.	Pol.	Ven.	Yug.
R^2	0.45	0.45	0.42	0.61	0.72	0.60	0.62
σ	0.11	0.12	0.09	0.09	0.07	0.09	0.08
Autocorrelation	0.75	0.74	0.58	2.20	1.01	1.53	0.75
Normality $X^2(2)$	1.34	0.89	0.33	1.55	17.17	0.05	7.38
Linearity $X^2(4)$	3.10	2.99	2.56	0.81	3.34	1.38	1.61
Homoscedasticity $X^2(1)$	2.30	0.41	12.30	0.00	0.18	5.25	0.21
Parameter Stability $X^2(9)$	14.43	5.78	16.15	22.09	11.46	8.60	16.53

Notes: (1) Figures in brackets are t-statistics.
(2) The tests for autocorrelation are distributed F(4,(59-k)), where k is the number of explanatory variables.
(3) OTHDUM1 and OTHDUM2 are as follows;

	OTHDUM1	OTHDUM2
Arg.	1990.5	1990.12
Bra.	1987.3	1990.3
Chi.	1991.4	
Mex.	Δ BRADYMEX	
Pol.	1991.3	
Ven.	1990.9	
Yug.	1987.9	1991.4

As regards the other regressors we found that the BRADYT dummy was significant in 5 of the countries examined. We also found that for most of the models there was the occasional month during which prices moved inexplicably fast requiring additional dummies to be used.[1] For the case of Argentina there were two such months - 1990.5 and 1990.12. The former appears to be associated with speculation that the IMF would reactivate its letter of intent since it and the Argentine government were felt to be close to reaching agreement. December 1990 witnessed the IMF approving Argentina's economic plan and deciding to resume disbursements on a standby credit (see the appropriate IFRs for details). Dummy variables were defined accordingly and used in our preferred model. Although at first blush the Argentine model may look a little sparse, it is worth noting that, once one substitutes for the lagged residuals term (using the preferred model from table 3), one finds that it includes a role for each of the potential explanatory variables tried - interest rates, "A", "B" and "C" factors (plus, of course, the dummies). This is a feature of most of the preferred models - all of the matrix factors turn out to play some role with the exception of interest rates in the Yugoslav model and "C" factors in the Venezuelan one.

As regards Brazil we used a dummy for 1987.3 because of Citibank's announcement that it was thinking of downgrading most of its Brazilian debt by placing it on a "cash only basis" (see Kemna (1990)). A second dummy, taking the value 1 in March 1990, was designed to pick up the price "crash" (in Brazilian debt) which occurred in the first week of the month following reports that the new central bank head had reservations about continuing with debt-debt or debt-equity conversions (see IFR of March 10 for details). Dropping these dummies makes little difference to the other estimated coefficients in the model.

The model of Chilean prices includes a lagged dependent variable with a coefficient above one-half, suggesting that prices in this market respond less fast to "news" than in the other countries considered. Despite this the model appears to forecast accurately (though with a slight tendency to underpredict the recent improvement in prices).

Looking at the Mexican model one notices that the "forecast" test result is failed indicating parameter instability. This is because the model is able to explain developments after 1990.7 less well than prior to then. In part this may simply reflect data problems - from 1990.4 onwards Mexican debt prices published by Salomons have referred to par and discount bonds resulting from the Brady package. (The equivalent debt to the old pre-Brady debt is no longer

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1 "Requiring" in the sense that otherwise tests for normality of the residuals tends to be failed (because of skewness and/or kurtosis). Note, however, that the addition of the dummies makes very little difference to the other estimated coefficients or to the models' overall fit.

traded in significant amounts and so prices for it are not available from Salomons.) We have had to use Merrill Lynch data for the period April to July and Chartered West LB data thereafter (both of which relate to the old base debt). These show a sharp rise in the price of Mexican debt in July. We have tried weighting together the prices of the par and discount bonds (according to the amount outstanding of each), but it is clear that there is still something of a structural break in the data. Although we have included dummies to permit both a temporary and permanent rise in prices because of this, the fact that the Mexican model fails the forecast test provides additional evidence for the view that Mexico's Brady deal has been viewed positively by creditors.

As regards Poland the model includes a (temporary) positive role for "B" factors, which is perhaps surprising (though inclusion of such a term is not crucial to the model's main features/fit). One possible explanation is that arrears build-ups have been viewed by creditors as signalling the closeness of Paris Club deals (which could improve prospects of the banks receiving some payments).

The Venezuelan model has the most significant (and largest long-run) effect from interest rates. This may explain why the economic ("C") factors did not turn out to be significant in explaining price movements (there is some double-counting of interest rate effects). Certainly the two are highly correlated (90% of the variation in interest rates being explained by the "C" factor scores).

Finally, as regards Yugoslavia, the estimated model turned out much better than expected given the doubts that we had found a cointegrating regression when examining the long-run model. Note, however, that the lagged dependent variable does quite a large proportion of the work in explaining price movements. The two dummies are defined so as to allow the model to explain the two price falls - the more recent undoubtedly being associated with political uncertainty.

VI: Conclusions

In this paper we have tried to improve upon existing models of secondary market prices (specifically that of Cohen and Portes (1990)). This we have done by:

(i) extending the dataset to include more recent data,

(ii) improving the dataset by changing the weighting procedure used in calculating average prices,

(iii) changing the functional form of the equation estimated (to avoid problems of inconsistency between the theoretical and statistical models used),

(iv) using econometric techniques which emphasise the time-series properties of the variables (and which allow us to test whether we have a "sensible" long-run model),

(iv) adding to the list of explanatory variables used by including the Bank of England's matrix scores as a regressor,

Our preferred model appears to explain past data well and exhibit stable parameters (and so may forecast adequately). More importantly it suggests that secondary market prices reflect not just political but also economic factors, (and thus can be explained in terms of the underlying "fundamentals"). Our results support recent research, such as that by Anayiotos and de Pinies (1990), which suggests the secondary market is broadly efficient (in that prices generally reflect much of the relevant new (current and forward looking) information), but that the illiquidity of the market does mean that occasionally large transactions (such as Citibank's provisioning decision) can have a major effect on the market. This is not particularly surprising given that the value of secondary market transactions though rising is still less than 5 % per annum of the value of developing country debt.

APPENDIX 1

Using secondary market prices to predict debt repayment problems.

We assume that secondary market prices are useful in predicting debt problems, in which case, if the matrix is useful in helping to predict prices, then it will also be useful in predicting debt problems (see page 1). This appendix briefly summarises previous research which bears out this assumption.

Hajivassiliou (1989) examines directly the relationship between secondary market prices and debtor creditworthiness and, more specifically, whether secondary market discounts simply reflect past repayment problems or are able to anticipate future debt crises. (The nature of any relationship has important implications for the desirability of debt relief - if the secondary market discounts can successfully anticipate future repayment problems then partial debt relief may be beneficial for both creditors and debtors if it helps the latter avert anticipated problems.)

First, Hajivassiliou reviews some of the econometric models used to quantify creditworthiness. Each of probit, ordered probit and tobit models specify y^*_{it} (the (unobserved) propensity of country i to have a debt problem in period t) as a (linear) function of observable country characteristics, past history, world conditions and other factors;

$$y^*_{it} = x'_{it}\beta + \varepsilon_{it} \qquad (1)$$

where the x matrix contains the explanatory variables, β is the vector of parameters to be estimated and ε the associated error terms. Once y exceeds a threshold a debt problem is observed (if one wishes to permit there to be different degrees of severity of debt problems in the model one might wish to include a number of thresholds). All three models are estimated using panel data (covering 109 developing countries over the period 1970 to 1986), finding that past problems and economic variables provide a fairly good base on which to forecast repayment problems. Since these factors are generally observable to private creditors there seems little reason why secondary markets should not also reflect similar considerations. If they do so (and the markets are efficient) then secondary market prices could help explain changes in the creditworthiness indicators (used as the dependent variable in (1) above). This is what Hajivassiliou finds when the creditworthiness indicators are regressed on the secondary market price - the t-values on the latter are in nearly every case significant. (Allowance is made for persistent heterogeneity among nations and for the impact on bankers' perceptions of a history of bad debt performance.) The author admits, however, that the correlations are not particularly strong, suggesting that "secondary market evaluations anticipate only rather imperfectly future financing difficulties". This suggests that further research is needed if the assumption we made is to be recognised as having been a reasonable one to make.

APPENDIX 2

This appendix gives details of the Bank of England's matrix. This is offered as guidance to assist banks in the determination of country debt provisions.

There are three stages in the process of deciding an appropriate level of provision:

(i) to indentify countries with current or potential repayment difficulties;

(ii) to identify the nature of those difficulties and the extent of the country's problems; and

(iii) to determine, at this point, what proportion of exposures to that country is unlikely to be repaid in full.

Accordingly three categories of factors are used in the matrix:

A Factors - these attempt to gauge a borrower's inability or unwillingness to meet its obligations, whether at the due date or thereafter;

B Factors - these indicate a borrower's current difficulties in meeting its obligations; and

C Factors - these provide evidence of the likelihood of repayment difficulties either persisting or arising in the future.

Altogether sixteen factors are included. The attached table shows the individual factors and the scores used to weight them together.[1] Note that only one factor is weighted within a range according to individual judgement ("other factors"). Note also that one of the factors is secondary market prices. Obviously, the scores for this factor were subtracted from the totals in constructing the matrix scores to be used as a regressor in explaining prices.

When setting provisions a moving average of the matrix scores is used - each score being allocated to a provisioning band. In order to carry out our analysis we used the raw scores, rather than the moving average. Since we score the matrix only once a quarter, we interpolated the raw scores to obtain a run of monthly scores for use in our regression work.

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1 Further details are available from the authors.

Matrix factors and scores.

Factor:	Score:
Moratorium in effect.	Up to 12.
Rescheduled at any time in the last 5 years or in the process of rescheduling.	Up to 12.
Second or more rescheduling during the last 5 years of principal amounts rescheduled since January 1983.	Up to 12.
Significant arrears of interest or principal to IFIs.	Up to 15.
Arrears of principal on original or rescheduled loans from other external creditors.	Up to 10.
Arrears of interest on original or rescheduled loans from other external creditors.	Up to 10.
New money following rescheduling to clear arrears.	Up to 10.
Interest service ratio.	Up to 10.
Visible import cover.	Up to 10.
Debt-GDP ratio.	Up to 10.
Debt-exports ratio.	Up to 10.
Not meeting IMF targets.	3.
Unfilled financing gap over next 12 months.	2.
Secondary market price.	Up to 12.
Highly dependent on one source of income.	2.
Other factors.	Up to 5.
Total score:	**Up to 145.**

APPENDIX 3

Test results to find the order of integration of each of the variables used in both our general and country specific models are presented below. Both Dickey-Fuller (DF) and Augmented Dickey-Fuller (ADF) statistics are shown, the latter having permitted lags of up to 4 periods to be used in the test regression.

General models

Variable	Levels		Differences	
	DF	ADF	DF	ADF
ln S	-1.54	-1.62	-7.40	-2.80
ln r	-0.69	-1.60	-5.94	-2.91
ln TS	-1.79	-1.80	-6.79	-2.72
MS	0.16	0.98	-2.65	-3.87
MSA	0.07	-1.31	-3.27	-3.75
MSB	0.16	-1.05	-2.35	-3.72
MSC	-0.81	-1.42	-2.79	-4.96

Country-specific models

Country	Variable	Levels		Differences	
		DF	ADF	DF	ADF
Argentina	ln S	-1.61	-1.91	-7.44	-2.05
	MSA	0.31	-0.26	-2.81	-7.00
	MSB	-0.56	-0.94	-2.49	-4.65
	MSC	-0.74	-1.25	-3.47	-6.50
	MS	0.14	-0.78	-3.16	-4.49
Brazil	ln S	-1.49	-1.65	-6.97	-2.89
	MSA	-1.08	-1.65	-2.74	-4.09
	MSB	-0.51	-1.72	-2.55	-3.18
	MSC	-0.69	-1.53	-2.04	-3.71
	MS	-0.28	-1.40	-2.28	-3.75
Chile	ln S	2.27	1.23	-4.50	-2.44
	MSA	-1.34	-1.87	-2.78	-4.17
	MSB	-	-	-	-
	MSC	-0.76	-1.24	-2.62	-5.19
	MS	-0.54	-1.29	-2.31	-4.35

Mexico	ln S	-0.31	-0.42	-7.09	-2.22
	MSA	-0.88	-1.67	-2.29	-3.78
	MSB	-	-	-	-
	MSC	-1.78	-2.93	-3.36	-4.70
	MS	-1.43	-1.83	-3.13	-5.09
Poland	ln S	-1.46	-1.57	-4.70	-2.70
	MSA	-0.73	-1.17	-2.90	-5.17
	MSB	-1.34	-1.81	-3.48	-4.94
	MSC	-1.61	-2.17	-0.14	-3.48
	MS	-0.89	-2.17	-2.98	-3.92
Venezuela	ln S	-1.69	-1.66	-5.73	-2.42
	MSA	-0.14	-0.81	-2.74	-4.87
	MSB	-0.70	-2.25	-2.09	-4.76
	MSC	-1.87	-2.51	-2.49	-5.21
	MS	-1.19	-2.38	-2.80	-4.07
Yugoslav.	ln S	-1.07	-1.02	-5.67	-2.38
	MSA	-1.39	-2.25	-2.88	-5.35
	MSB	-0.94	-2.17	-2.61	-3.47
	MSC	-0.93	-2.24	-3.09*	-6.30*
	MS	-1.18	-2.19	-2.27	-3.66

* Second differences used (first differences gave values below 3 in absolute terms).

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