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Introduction

The article considers the problem of analyzing product price inter-relatedness to determine whether markets are competitive. This notion goes back to earlier analysis undertaken in the 70s (Maunder, 1972), but these studies pre-date the application of tests of stationarity and a well developed understanding of this issue in the economics literature. An interesting survey of the application of time series methods to explain pricing discrepancies and test for the existence of a causal relation that underlies pricing behaviour was prepared by LecG in 1999 for the United Kingdom Office of Fair Trade (OFT). Forni (2004) reported on a study that he prepared on milk pricing across Italy. It was suggested that tests under the alterative and the null of stationarity can be applied as an effective and efficient way of analyzing anti-trust cases associated with market concentration and non-competitive pricing. Tests of stationarity in the context of a price differential are essentially tests of cointegration between price series and cointegration is viewed in this sense as a signal of a broad market definition that is consistent with arbitrage (Forni, 2004). The cointegration case has been used to determine competitiveness in a number of studies, see De Vany and Walls (1999), Hendry and Juselius (2001) and La Cour and Møllgaard (2002).

In this article we discuss further the developments in the literature previously summarized in Hendry and Juselius (2001). We find that beyond the narrow regional and geographical arguments that apply to the study of Forni (2004), it would seem difficult to concede that stationarity tests applied to determine that price proportions are stationary can dominate the richer analysis that can be drawn from the cointegration approach. We feel that the approach that is developed from testing stationarity has some potential pitfalls that may not be recognized when all one is interested in is the behaviour of a ratio. As is explained by Hendry and Juselius, there are occasions when tests of stationarity may be more effective and more efficient, but they are conditioned on the prior observation that the model underlying this approach is valid.

Firstly we provide a brief summary of the literature on stationarity and error correction, but draw the reader's attention to the quite considerable literature on the use of statistical methods to assess anti-trust cases.[i] Then we consider a trivariate extension of the model analyzed by Hendry and Juselius (2001). Finally we develop an empirical example to motivate our analysis and then make our conclusions.

Stationarity tests and measures of market definition

The notion that an analysis of price behaviour might be used to detect irregularities in pricing can be found in studies such as Maunder (1972). As is correctly explained by Forni (2004), this discussion predates an appropriate understanding of the pitfalls that arise when inference is undertaken on time series that are non-stationary. Forni (2004) has addressed this issue by suggesting that some form of test of stationarity can be applied to logarithmic data on prices. Specifically, Forni applies his analysis to data on a homogenous product to determine whether there are local markets for milk in Italy. The analysis is undertaken on a relatively small data set, the bivariate comparisons are applied inter-regionally and they relate to homogenous products where local legislation places a limit on the capacity to sell the product easily across Italy. Forni, considers two forms of test and sets up a framework by which the combination of tests under the

(under the alternative) it is a test of the proposition that two or more series are stationary in combination or rather deducting part of one series from another renders the first series stationary. Firstly, we consider the problems that arise by not considering (1) above to be a model of either price, then we move on to discuss the approach adopted in De Vany and Walls (1999). The latter study realizes that the model estimated by the app be used to detect discriminatory or unfair pricing. In the PPP literature, Banerjee et al (2005) suggest that the single equation or panel approach are not always correct, because they may miss out on the important cross sectional dependence between price series, the exchange rate and other fundamental factors that can also influence the real exchange rate.[v] There is also some evidence from the empirical literature based on multivariate tests of cointegration that the proportionality restriction associated with PPP can be accepted when series are modelled jointly (Johansen and Juselius, 1992, and Juselius, 1995). The latter point seems very important for the validity of univariate analysis of relative price behaviour as unlike PPP there may be additional reasons for price movements not to match each other.

In the case of PPP it is important that the basket of goods be comparable. Lothian (2012) who analyses 400 years of price and exchange rate data for Great Britain and Holland provides evidence for the importance of the nature of the price series and compares a range of different historical price series across this period. The key issue in the context of pricing of prodfti04(e)-5.5(Bcli)e aries

These two conditions are necessary and sufficient for weak exogeneity of z_t for the sub-block $_{.1}$. It follows that weak exogeneity for a sub-block occurs when the condition $_{2,1}=0$ is extended to a quasi diagonal form:

$$\begin{matrix} {}_{1,1} & 0 \\ 0 & {}_{2,2} \end{matrix} .$$

Where $_{1,2} \quad {}_{1,2}(\;{}_{2,2})^{1}\;{}_{2,2} \quad 0$ or either $_{2,2}$ or $\quad {}_{1,2}(\;{}_{2,2})^{1} \quad 0$. Otherwise $_{1,2} \quad {}_{2,2}$ and:

$$\begin{matrix} {}_{1,1} & {}_{2,2} \\ 0 & {}_{2,2} \end{matrix} .$$

The loadings in the first and second blocks are to proportional to each other. In the case where $_{1,2} \quad 0$, then weak exogeneity of z_t for the matrix occurs when $_{2,2} \quad 0$ and:

$$\begin{matrix} {}_{1,1} & {}_{1,2} \\ {}_{2,1} & {}_{2,2} \end{matrix} \quad \begin{matrix} {}_{1,1} & {}_{1,2} \\ 0 & 0 \end{matrix} \quad \begin{matrix} {}_{1,1} & {}_{2,1} \\ {}_{2,1} & {}_{2,2} \end{matrix}$$

$$\begin{matrix} {}_{1,1}\;{}_{1,1} \quad {}_{1,2}\;{}_{2,1} & {}_{1,1}\;{}_{2,1} \quad {}_{1,2}\;{}_{2,2} \\ 0 & 0 \end{matrix} .$$

As a result the cointegrating vectors are estimated from the block of equations related to *y*. Notice that there are at most *N*-*r* weakly exogenous variables in this sense or the fundamental criterion for cointegration is broken.

Cointegrating exogeneity augments the triangular with non-causality between *y* and *z* at the level of the system. Hence, the long-run relationships for *z* do not depend on the levels of *y*. It follows that z_t is cointegrating exogenous for the sub-vector $_{.1}$, if and only if:

$$_{2,1} \quad 0.$$

$_{2t} \quad {}_{2,2}z_t$ define r_1 and r_2 blocks of stationary variables.

If conditions (i) and (ii) above hold, then cointegrating exogeneity in this form is an exact analogue of strong exogeneity (see Engle et al, 1983).

$$A \quad x_t \quad \begin{matrix} 1 & 1 \\ 0 & 1 \end{matrix} \quad \begin{matrix} 11 \\ 12 \end{matrix} \; [1 \quad {}_{12}] x_{t \;\; 1}$$

$$\begin{matrix} 1 & 1 \\ 0 & 1 \end{matrix} \quad \begin{matrix} 11 & 12 \\ 21 & 22 \end{matrix} \quad x_{t \;\; i} \quad \begin{matrix} 1 & 1 \\ 0 & 1 \end{matrix} \quad {}_t$$

$$\begin{matrix} p_{1t} \quad p_{2t} \\ p_{2t} \end{matrix} \quad \begin{matrix} 11 & 12 \\ & 12 \end{matrix} \; [1 \quad {}_{12}] x_{t \;\; 1}$$

$$\begin{matrix} 11 & 21 & 12 & 22 \\ & 21 & & 22 \end{matrix} \quad x_{t \;\; i} \quad \begin{matrix} 1t & 2t \\ & 2t \end{matrix} \; .$$

Notice, that the transformation makes no difference to the long-run equations. For inference in the single equation to be equivalent to inference that arises in the VAR we require the following restrictions to hold, ${}_{12}$ ${}_{22}$ 0, ${}_{12}$ 0 and ${}_{12}$ 1:

$$\begin{matrix} p_{1t} \quad p_{2t} \\ p_{2t} \end{matrix} \quad \begin{matrix} 11 \\ 0 \end{matrix} \; [1 \quad {}_{12}] x_{t \;\; 1}$$

$$\begin{matrix} 11 & 21 & 0 \\ & 21 & 22 \end{matrix} \quad x_{t \;\; 1} \quad \begin{matrix} 1t & 2t \\ & 2t \end{matrix} \; .$$

In the case of Forni (2004), the Dickey-Fuller test imposes the restriction ${}_{12}$ 1on the long-run and also imposes the same restrictions on the short-run dynamics of an error correction model that contains no cross price terms or ${}_{12}$ ${}_{22}$. When the VAR is compared with the single equation error correction framework then inference is optimal in the long-run as this is efficiently estimated on a single equation when p_{2t} is weakly exogenous for the cointl8 Tc9ekm.g86.480djET.4(e)6-4.5((-run)(tl8 Tc4.

$$\Delta\begin{bmatrix} p_{1t} \\ p_{2t} \\ p_{3t} \end{bmatrix} \quad \begin{bmatrix} {}_{11} & {}_{12} \\ {}_{21} & {}_{22} \\ {}_{31} & {}_{32} \end{bmatrix}\begin{bmatrix} 1 & {}_{21} & {}_{31} \\ {}_{12} & 1 & {}_{32} \end{bmatrix}\begin{bmatrix} p_{1t?1} \\ p_{2t?1} \\ p_{3t?1} \end{bmatrix}$$

$$?\begin{bmatrix} {}_{11} & {}_{12} & {}_{13} \\ {}_{21} & {}_{22} & {}_{23} \\ {}_{31} & {}_{32} & {}_{33} \end{bmatrix}$$

Setting the initial condition to zero and for the purposes of exposition =0, then the common trend can be characterized by:

$$\acute{o}' x_t \qquad \acute{o}' \sum_{i=1}^{t} \quad {}_i .$$

Therefore with N-r=1 a single common trend drives prices. It follows from the Granger representation theorem that this single common trend is annihilated by the cointegrating vectors that define r stationary variables when the condition for cointegration $C \quad 0$ *is satisfied. To see this multiply (4) by :*

$$\begin{aligned} {}' x_t \qquad {}' C x_0 \qquad {}' C\Big(\sum_{i=1}^{t} \quad {}_i\Big) \qquad {}' \;(\quad {}' \quad)^{?1} \sum_{i=0}^{K} (I \qquad {}' \quad)^{i} \; {}'(\quad {}_i) \\ \sum_{i=0}^{K} (I \qquad {}' \quad)^{i} \; {}'(\quad {}_i). \end{aligned}$$

This defines r stationary series without restriction on the cointegrating relationships. The stationary variables, x_t *are linear combinations of the prices that return the system to equilibrium. For a broad market definition we require the pri,-.t8o move inTw(e46.5(Tf.3722 0 TD(.)Tj/TT4 1 Tf1*

is of interest to note that economic theory implies that all firms ought to react to the underlying competitive behaviour, that is per-force associated with a common trend when the series are all

$$\Delta\begin{bmatrix} p_{1t} \\ p_{2t} \\ p_{3t} \end{bmatrix} \quad \begin{bmatrix} 1 & a_{12} & a_{13} \\ & & \\ & & \end{bmatrix}$$

tested prior to this re-arrangement.

The hypothesis $p_{1t} \quad {}_{12}p_{2t} \quad I(0)$ may not give rise to a unit coefficient, because the prices have a different response or the prices relate to variables that are not homogenous or are differentiated in some way or other (Hosken and Taylor, 2004).

There might be other prices or other variables, which might legitimately influence the long-run relationship.[x] In the context of PPP, interest rates are often required to make the real exchange rate stationary (Johansen and Juselius, 1992 and Juselius, 1995).

Empirical Analysis

Here we consider similar data to Hosken and Taylor (2004), Hunter and Tabaghdehi (2012) and Kurita (2008) to analyze the stationarity and cointegration properties of regional gasoline prices in the US. The data is similar to that analyzed by Kurita (2008) who took the same regions for the period 1990 (week 38) - 2004 (week 52) and this is a subset of the data analyzed by Hunter and Tabaghdehi (2012) who consider a broader range of regions. Here the price series are tested for stationarity and these results are compared with those related to the Johansen procedure. The data are for three regions of the US and the product is a homogenous product (gasoline) except for the location. Here three regions are selected, New York (NY), the Gulf Coast (GC) and Lower Atlantic (LA). The regions are similar, they are all coastal regions, they each had facilities by which crude oil can be imported and refined.

In Figure 1 the logarithms of gasoline prices are given for the data used by Kurita (2008) who considers three prices from the regions, New York (NY), Gulf Coast (GC) and Lower Atlantic (LA):

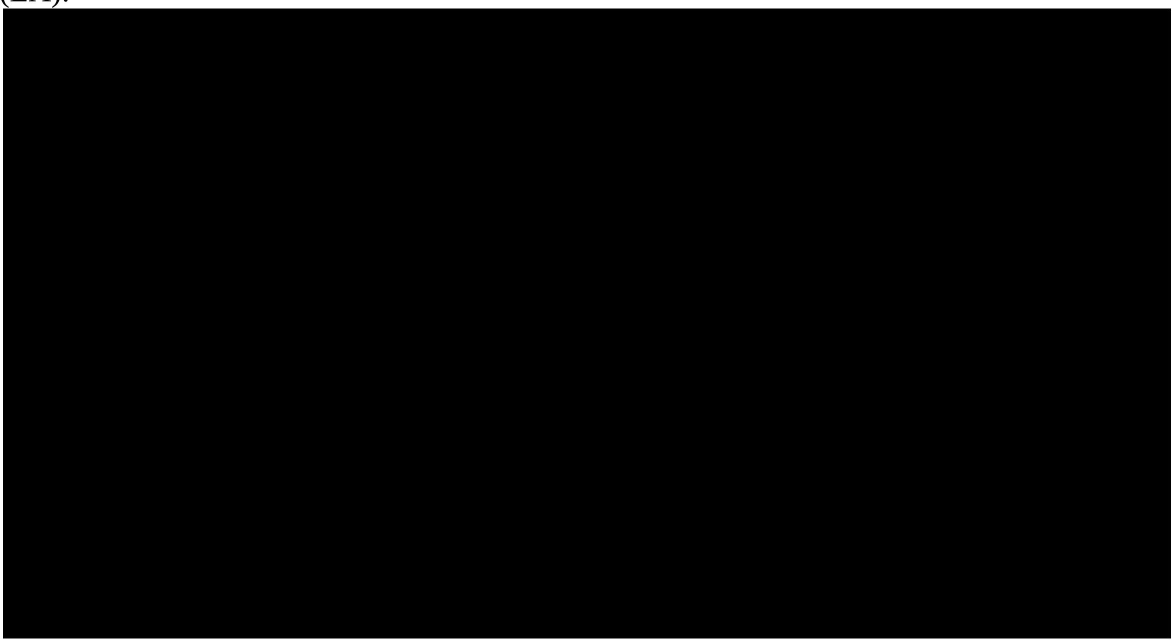

Figure 1: The logarithm of gasoline prices for NY, GC and LA

In Figure 2, the series have also been transformed into log price differentials, so as is explained in section 2, the notion of cointegration is tested using Dickey Fuller regressions (Forni, 2004) and Error Correction Models (Kremer et al, 1992).

Figure 2: Logarithmic price differentials: NY and LA, GC and LA, GC and NY.

observe that the results in the first and the final element of this column are the same as the latter is the reverse regression. This suggests that the method developed in Forni (2004) should be limited to either the upper or lower triangle of results. Hunter and Tabaghdehi (2012) show that there is an exact correspondence to this result for the Dickey Fuller (DF) test and the estimations in the first column suggest that this generalizes when *q* lags appear in the ADF test estimations.

There are instances where it may be appropriate to test using the ADF test. This arises either when the sample is small or when the price proportions exhibit less complex time series behaviour. If there are common features in the data that cancel then the stationarity tests may reveal an effective methodology to determine the nature of the market. As can be observed from Table 1 the dynamic in some cases is restricted to 11 and 17 lags, this compares to the VAR that is estimated with 20 difference terms. However, the approach is limited to finding whether prices move in proportion and is not able to determine anything about causality or conditioning. Further, the autoregressive (AR) model that underlies the ADF test imposes a common factor restriction on the ECM (Burke and Hunter, 2005).

The test is efficient when the model can be restricted so that the prices move in proportion in both the long and the short-run. This restriction can be removed by augmenting the DF test with either of the price inflation terms. The second column in Table 1 is based on the model underlying the ADF test extended by the contemporaneous inflation terms from the second price in each of the correction terms (p_{it}) and their *q* lags. For conventional inference to be efficient, the augmenting price inflation variable is required to be weakly exogenous. However, this is not a problem for the test of stationarity on the error correction term as the estimated coefficient is super consistent (Davidson and MacKinnon, 2004). Here, these tests are not enhanced by augmentation with the price inflation terms, though it cannot be concluded from this that the market is also efficient in a variable, but this cannot be determined by directly estimating one equation and for the equation to be well defined, the cointegrating rank should also be determined in advance. As a result further discussion of the error correction model is considered in relation to the simulated data as the nature of the model is known in that case.

The system estimated is a trivariate unrestricted VAR(21) in error correction form. It can be observed from the first row of results in Table 2 that the null of the Johansen trace test can be rejected at both the 1% and 5% level. The subsequent element in the trace test relates to cointegrating rank of *r=3*, but as the test is not significant (p.value=[0.281]) so it follows that the $rank(\;)\quad N\quad 1\quad 2$. To preclude the possibility of I(2) trends it makes sense to undertake a test that will confirm the rank result (Johansen, 1995, and Burke and Hunter, 2005). This is a rank test for the null of non-stationarity with respect of the second differences and this is a sufficient condition for cointegration. If the I(2) test related to Paruolo (1996) is applied, then with *r=2* cointegrating vectors and *N=3* variables, here the key test for I(2) relates to the final statistic on the diagonal of the I(2) test table the test term is ${}_1Q_{2,1}\quad 49.586[0.0000]$. The clear significance of this test implies that there is a single stochastic trend and with *r=2* and so WE is rejected at the 1% level of significance. In the next block long-run exclusion (LE) is tested (Juselius, 1995). The clear significance of these tests emphasizes the rank condition and the likely robustness of inference associated with any normalization of either cointegrating vectors. The equations are ordered in relation to the WE test as is suggested in Hunter and Simpson (1996). WE is tested for LA on a model normalized on the price for NY in the first equation and the GC in the second equation.

Identification requires there to be *r-1* restrictions per long-run relation (Hunter and Simpson, 1996). This can be undertaken in a number of ways, but one that is straightforward relates to what is termed the normalization rule by Boswijk (1996). Therefore $I_2 \quad b$ where I_2 is the $2 \quad 2$ identity matrix and $b \quad b_1 \quad b_2$. This restriction is not binding as it exactly identifies each cointegrating vector so as might be observed for the next line in the table the WE test is not affected by applying the normalization rule to . The next row tests for CE by imposing the restriction $_{2,1} \quad 0$. This gives rise to a triangular block in so the long-run equations for GC and LA are not long-run caused by the NY price. However, this is seen in Hunter and Simpson (1996) as trivial in terms of the first long-run relation as the GC price does not appear in this vector.[xi] In the final block the notion of parallel pricing is tested by imposing the restrictions $b \quad 1 \quad 1$. As a result these tests give rise to the same conclusions as the stationarity tests.

As Forni (2004) correctly points out the test based on the error correction term simultaneously tests parallel pricing along with cointegration. However, it does not consider WE and CE or the orientation of the long-run. These issues are considered in more detail when the simulations are analyzed. There is further discussion of exogeneity and identification for this type of model in Burke and Hunter (2011).

The final set of results in Table 2 are accepted at the 5% level and they yield the following and matrices, and the associated restricted matrix conditioned on the GC price that is cointegrating exogenous and the LA price that is weakly exogenous for :

$$\begin{bmatrix} \underset{(0.035941)}{?0.18686} & \underset{(0.040008)}{?0.11187} \\ 0 & \underset{(0.021095)}{?0.079938} \\ 0 & 0 \end{bmatrix} \text{and} \quad ' \begin{bmatrix} 1 & 0 & ?1 \\ 0 & 1 & ?1 \end{bmatrix}$$

long-run relations seen as arbitrage correction terms. The results in Table 2 tell us that these responses are observed to flow in the direction LA to the other regions and this would suggest that the market is not informationally efficient (Hunter and Burke, 2007, and Kurita, 2008). This analysis would be strengthened were we to have information on prices at the firm level, but further information might be extracted from regional details on the supply of gasoline. Further, these results imply that the regulatory authorities would need to be careful to permit any further concentration in the industry especially in terms of gas station ownership or refinery capacity in LA. Similarly, a further causal feature relates to the GC price being cointegrating exogenous for $_{.1}$ 1 0 1 the first cointegrating vector.

Next the simulated data are considered and the following matrix imposes WE and is fixed so that there is parallel pricing. The same cointegrating vectors apply in the simulated case as occurs with the actual data:

$$ {}_{s} \begin{bmatrix} ?.09 & .04 \\ 0.11 & ?.17 \\ 0 & 0 \end{bmatrix} \text{and} \quad {}'_{s} \begin{bmatrix} 1 & 0 & ?1 \\ 0 & 1 & ?1 \end{bmatrix}. $$

though it is based on a simpler time series structure.[xii] The series are computed as a system, but the table below yields single equation results that follow from the simulated data. It is of interest that based on simulated data Haug (1996) found that there was no correlation between the Johansen test and the residual based cointegration tests. However, the simulations in Haug were generated under the null while the results presented in Table 3 relate to a cointegration case.

Table 3, Unit root and ECM tests of price-proportions for simulated data.

Correction terms	$\hat{\ }_{1}$ 1	t ? DF	t ? ecm
$p_{1?3}$	0.89927	-5.974 DD	-8.09 DD
$p_{2?3}$	0.92059	-5.247 [redacted]	-9.15 DD

two elements in the last column. The next p.value in the column is associated with a small test statistic and this implies insignificance; this arises as the third variable is weakly exogenous for . By sequentially estimating the error correction models it is possible to determine whether a series

Conclusion

There is a simplicity associated with the method that applies tests of stationarity to the problem (Forni, 2004). However, as is shown in Hunter and Tabaghdehi (2012) behind the simplicity are a number of problems. In particular, testing the reverse results and repeated testing when it is known that there are *N-1* cointegrating vectors at most. Then the common factor restriction inherent in the DF/ADF test binds the same behaviour to the short-run and the long-run. However, when the time series is not extensive then it may be that there is no alternative that make sense and it is also possible to extend the approach to the panel context (Beirne et al, 2007, Hadri, 2000, and Giulietti et al, 2010). Error correction models may have a similar felicity of application and possibility of extension to the panel case and they also have some further benefits in regard to testing (Kremers et al, 1992).

However, stationarity might have been rejected in the context of DF equations, because: the tests perform poorly with relatively small samples (Podivinsky and King, 2000); the test has a less appropriate null (Hosken and Taylor, 2004) or the proportionality restriction required to test stationarity as compared with cointegration is too restrictive (Hunter, 2003). Key variables may also be omitted from the long-run so the dynamic equations do not meet the conditions required for separate estimation and inference (Hunter and Burke, 2007 and Kurita, 2008). The meitaf3-.8(er-15.3(itaf0l07 Tr) that the GC can in isolation be seen as weakly exogenous for , but subject to LA being weakly exogenous, it is then only possible for GC prices to be cointegrating exogenous for $_{.1}$. This implies that the long-run relations associated with GC prices are not reflecting what happens in NY and this is a further signal of imperfection. It should be noted that finding a system with *N-1* long-run relations all satisfying the pure error correction terms related to parallel prices is also consistent with what is termed Long-run Equilibrium Price Targeting by Burke and Hunter (2011).

References

Banerjee, A., Marcellino, M. and Osbat C., (2005) Testing for PPP: Should we use Panel Methods? Empirical Economics, 30, pp. 77-91.

Bauwens, L., and Hunter, J. (2000) Identifying long-run behaviour with non-stationary data. Discussion Paper 2000/43 CORE, The Catholic University, Louvain-la-Nueve.

Beirne, J., Hunter, J., and Simpson, M., (2007), `Is the Real Exchange Rate Stationary? The Application of Similar Tests for a Unit Root in the Univariate and Panel Cases', Qualitative and Quantitative Analysis in the Social Sciences, 1 (2), pp. 55-70.

Boswijk, H.P. (1992) Cointegration, Identification and Exogeneity: Inference in Structural Error Correction Models. Amsterdam: Thesis Publishers, Amsterdam.

Boswijk, H.P., (1996) `Cointegration, identification and exogeneity: inference in structural error correction models', Journal of Business and Economic Statistics, 14, pp. 153-160.

Burke, S.P. and Hunter, J., (2005) Modelling Non-stationary Economic Time Series: A Multivariate Approach. Palgrave, Basingstoke.

Burke, S.P. and Hunter, J., (2011) `Long-run, Equilibrium Price Targeting', Qualitative and Quantitative Analysis in the Social Sciences, 5 (1), pp. 26-36.

Caner, M. and Kilian, L., (2001) `Size Distortions of Tests of the Null Hypothesis of Stationarity: Evidence and Implications for the PPP Debate'. Journal of International Money and Finance, 20, pp. 639-657.

Davidson, J.E.H., Hendry, D.F., Srba, F. and Yeo, S., (1978) `Econometric modelling of the aggregate time series relationships between consumers expenditure and income in the United Kingdom', Economic Journal, 88, pp. 661-92.

Davidson, R. and MacKinnon, J.G., (1993) Estimation and Inference in Econometrics. Oxford University Press, New York.

Davidson, R., and MacKinnon, J.G., (2004) Econometric Theory and Methods. Oxford University Press, Oxford.

De Vany, A.S. and Walls, D., (1999) `Cointegration analysis of spot electricity prices: insights on

estimation and testing', Econometrica, 55, pp. 251-276.
Ericsson, N.R. and Irons, J.S., (1994) Testing Exogeneity. Oxford University Press, Oxford.
Forni, M., (2004) `Using stationarity tests in antitrust market definition', American Law and Economics Review, 6, pp. 441-464.
Froeb, L.M. and Werden, G.J., (1998) `A robust test for consumer welfare enhancing mergers among sellers of homogeneous products', *Economics Letters*, 58, pp. 367-369.
Giannini, C. and Mosconi, R., (1992) `Non-causality in cointegrated systems: representation, estimation and testing', Oxford Bulletin of Economics and Statistics, 54, pp. 399-417.
Giulietti, M., Otero, J. and Waterson, M., (2010) `Pricing behaviour under competition in the UK electricity supply industry', Oxford Economic Papers, 62, pp. 478-503.
Hadri, K. (2000) `Testing the Null Hypothesis of Stationarity against the Alternative of a Unit Root in Panel Data with Serially Correlated Errors', The Econometrics Journal, 3, pp. 148-161.
Haug, A.A., (1996) `Tests for cointegration: A Monte Carlo comparison', Journal of Econometrics, 71, pp. 89-115.
Hendry, D.F. and Juselius, K., (2001) `Explaining cointegration analysis: Part II', The Energy Journal, 22, pp. 75-120.
Hosken, D., O'Brien, D., Vita, M. and Scheffman, D., (2002) Demand Systems Estimation and Its Application to Horizontal Merger Analysis. Working paper 246, Federal Trade Commission Bureau of Economics, Washington.
Hosken, D. and Taylor, C.T., (2004) `Discussion of: Using stationarity tests in antitrust market

Juselius, K., (1995) `Do Purchasing Power Parity and Uncovered Interest Rate Parity Hold in the Long-run? - An Example of Likelihood Inference in a Multivariate Time-Series Model', Journal of Econometrics, 69, pp. 178-210.
Kremers, J.J.M., Ericsson, N.R. and Dolado, J., (1992) `The power of cointegration tests', Oxford Bulletin of Economics and Statistics. 54, pp. 325-348.
Kurita, T., (2008) Common stochastic trends and long-run price leadership in the US gasoline market. Working paper WP-2008-003. Faculty of Economics, Fukuoka University, Fukuoka.
La Cour, L.F. and Møllgaard, H.P., (2002) `Market Domination: Tests Applied to the Danish Cement Industry', European Journal of Law and Economics, 14, pp. 99-127.
LecG Ltd., (1999), Quantitative techniques in competition analysis. Research paper 17, OFT 266, Office of Fair Trading, London.
Leybourne, S.J. and McCabe, B.M.P., (1994) `A Consistent test for a Unit Root', Journal of Business and Economic Statistics, 12, pp. 157-166.
London Economics., (2002). Study of Sectoral Price Inter-dependence. Prepared for the Nederlandse Mededingingsautoriteit, Amsterdam.
Lothian, J., (2012) Exchange Rates and Prices in the Netherlands and Britain over the Past Four Centuries. Presented at the Money Macro Finance conference, Empirical Modelling of Financial Markets, Brunel University, May 23*rd* 2012, London.
Markham, J., (1951) `The Nature and Significance of Price Leadership', American Economic Review, 41, pp. 891-905.
Maunder, P., (1972) `Price Leadership an Appraisal of its Quality in some British Industries', The Business Economist, 4, pp. 132-140.
Paruolo, P., (1996) `On the determination of integration indices in I(2) systems', Journal of Econometrics, 72, pp. 313-56.
Podivinsky, J.M. and King, M.L. (2000) The Exact Power Envelope of Tests for a Unit Root. University of Southampton Discussion Papers in Economics and Econometrics, No. 0026, Southampton.
Sekioua, S. and Karanasos, M., (2006) `The real exchange rate and the purchasing power parity puzzle: Further evidence', Applied Financial Economics, 16, pp. 199-211.
Smith, R.P. and Hunter, J., (1985) `Cross Arbitrage and Specification in Exchange Rate Models', Economics Letters 18, pp. 375-376.
Toda, H.Y. and Phillips, P.C.B., (1994) `Vector autoregression and causality: a theoretical overview and simulation study', Econometric Reviews, 13, pp. 259-85.

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[i] See the discussion by Forni (2004) of the residual demand approach and any analysis that draws on the calculation of elasticities. Hosken, O'Brien, Scheffman and Vita (2002) consider demand studies based on systems to analyze the appropriateness of horizontal mergers. And Froeb and Werden (1998) discuss the use of the difference in the Herfindahl Hirschman index to assess welfare gains through mergers in homogenous product markets. There is also discussion of consumer surplus and following that consumer detriment (Hunter et al, 2001).

[ii] The Dutch regulator (Nederlandse Mededingingsautoriteit) has applied tests of non-stationarity to challenge mobile service providers (KPN, Orange and Vodafone) over the competitiveness of their pricing. However, the report by London Economics that analyses the Dutch market has been strongly criticized by Hunter (2003). The methods applied were incoherent and the data not sufficient to prove the case.

[iii] Forni (2004) makes the valid point that following Engle and Granger (1987) the series ought to be of the same order of integration for the test to be appropriately applied, but this suggestion gives the approach a broader appeal. So it should be anticipated that these series ought to be similar when they relate to an appropriately defined market. Furthermore, in the context of I(2) series as is explained in Johansen (1995) and Burke and Hunter (2005) the stationary combination may exist but not in the form considered by Engle and Granger (1987). Hence, the broad market definition may be wrongly rejected, because the wrong estimator

is applied. Similarly, the series may be fractionally integrated and again cointegration may be wrongly rejected when the I(1) approach is incorrectly thought to provide an acceptable approximation.

[iv] It should be noted that Forni (2004), cited by London Economics, applies both the ADF and the KPSS test, which uses the null of stationarity. When stationarity is an issue both of these tests should be considered. However, the stationarity test due to Leybourne and McCabe (1994) and the KPSS test used by Forni (2004) came under significant scrutiny by Caner and Killian (2001) who suggest that the tests under the null of stationarity perform relatively poorly. Sekioua and Karanasos (2006) provide evidence for the superior performance of the Generalised Least Squares Dickey Fuller test.

[v] Beirne et al (2007) suggest for the case of the real exchange rate that the relevant concept is that a sequence of exchange rates or relative prices are stationary on average. This leads to a tension between applying univariate tests where one may concentrate on specification of an underlying time series model and define corrected specifications, as compared with the panel approach where key aspects of the specification are assumed to be homogeneous. For the panel estimations presented in Beirne et al (2007) they argue when the test used applies the null of stationarity that the test due to Hadri (2000) works well with excellent power and size properties when the time dimension exceeds 50. It should be noted that Giulietti et al (2010) have had less success in this context though the electricity price data they use is more likely to be sensitive to non-linearity.

[vi] More generally, the restrictions can be imposed and the likelihood re-evaluated to test propositions on the cointegrating vect