Research Program on Forecasting

Liquidity effects on consumers' imports in Trinidad and Tobago

Michael Browne
The George Washington University
mbrowne@gwu.edu

RPF Working Paper No. 2016-005
http://www.gwu.edu/~forcpgm/2016-005.pdf

May 16, 2016

RESEARCH PROGRAM ON FORECASTING
Center of Economic Research
Department of Economics
The George Washington University
Washington, DC 20052
http://www.gwu.edu/~forcpgm

Liquidity effects on consumers' imports in Trinidad and Tobago*

Michael Browne†

May 16, 2016

Abstract

This paper examines the effects of liquidity on the demand for imports of non-durable consumers' goods in Trinidad and Tobago. A parsimonious vector equilibrium correction model (VEqCM) is used to test the hypotheses that liquidity has both long- and short-run effects. The multivariate cointegration approach of Johansen and Juselius (1990) is used to determine long-run relations and general to specific (*Gets*) modeling, *Autometrics* in particular, to determine system dynamic specification. Cointegration analysis reveals a long-run relation among consumers' imports, output, liquidity and relative prices. *Gets* modeling also reveals significant short-run liquidity effects and furthermore asymmetric short-run foreign and domestic price effects. The VEqCM seems a sufficient approximation of the underlying data generation process, demonstrates desirable statistical properties and empirically constant parameters, encompasses the findings of previous specifications and produces reliable forecasts.

JEL classification: F10; F14; F40; F41

Keywords: Trinidad and Tobago; imports; non-durable consumers' goods; liquidity; error correction model; cointegration; exogeneity; systems modeling; systems forecasts

*The author would like to thank Professors Neil Ericsson, Fred Joutz, Tara Sinclair and Michael Bradley, and participants at the 17th OxMetrics User Conference in Washington, DC, 2016 for invaluable comments and suggestions in the writing of this paper. Any errors are solely the author's responsibility. All numerical results were obtained using PcGive Version 14.0 and Autometrics Version 1.5e in OxMetrics Version 7: see Doornik and Hendry (2013).

†George Washington University Ph.D candidate. mbrowne@gwu.edu.

1 Introduction

Several studies in the consumers' expenditure literature have documented significant liquidity effects on aggregate consumers' expenditure, see for example Davidson *et al.* (1978), Hendry and von Ungern-Sternberg (1981) and Campos and Ericsson (1988). Despite this, studies on consumers' expenditure on imports have nonetheless abstracted from examining liquidity effects. This fact is surprising given that in small open economies, such as the small island developing states (SIDS) of the Caribbean and small non-renewable natural resource economies (SNRNRE), a large share of consumers' expenditure is likely to be imported. Given this potentially significant role of consumers' expenditure on imports in aggregate consumers' expenditure and the well-documented effects of liquidity on the latter, it is plausible that liquidity also has significant effects on consumers' expenditure on imports. These effects therefore appear natural hypotheses to be tested.

My research thus foremost examines liquidity effects, measured as start of period M2 money supply, on the demand for imports of non-durable consumers' goods, henceforth consumers' imports. The case of Trinidad and Tobago (TT) is examined as it is ideally categorized as both a SIDS and a SNRNRE. A parsimonious vector equilibrium correction model (VEqCM) is developed for this analysis and is used to test the hypotheses that liquidity has both long- and short-run effects on consumers' imports. The VEqCM is also used to test whether there are asymmetric foreign and domestic price effects, strong and/or super exogenous dependent variables, the ability of the system specification to account for the results of other models and the forecast accuracy of the system.

My research makes several contributions to consumers' expenditure on imports literature. First, it identifies significant long- and short-run liquidity effects on consumers' imports. Notably, cointegration analysis reveals a long-run relation among consumers' imports, the deviation of output from trend, liquidity and the ratio of foreign relative to domestic prices wherein liquidity serves as an integral control mechanism. Second, it identifies asymmetric short-run foreign and domestic price effects on consumers' imports. Specifically, changes in foreign prices are estimated to have an effect near three times that of changes in domestic prices. Third, it provides elasticity estimates for an updated TT quarterly dataset spanning the post-liberalization flexible exchange rate era. These elasticity estimates are generally in line with previous findings where applicable, with some notable caveats. Fourth, analysis reveals that long-run weak exogeneity of relative prices is rejected for consumers' imports; and that strong exogeneity but not super exogeneity is rejected for income and liquidity. Last but not least, the specification that incorporates liquidity effects is found to parameter encompass previous specifications and thus contributes notable research progression.

Empirical import demand models have been developed in several studies for TT and the greater Caribbean region, see for instance Watson and Teeucksingh (1997 and 2000), Modeste (2011), Hilaire *et al.* (1990), Watson (1990) and Watson and Clarke (1997). Generally, these studies have focused on modeling aggregate imports as a single endogenous equation in small macroeconometric models and were written during the period when the Central Bank of Trinidad and Tobago (CBTT) maintained a fixed exchange rate regime and before trade and financial liberalization led to TT becoming an economy characterized by high trade openness.[1] The fixed exchange rate would have, among other things, greatly mitigated exchange rate risk faced by domestic agents engaged in foreign trade, particularly on the import side considering a notable share of TT exports were petroleum products traded in US dollars. The *more* closed economy setting, however, would have significantly limited foreign trade. The policy decisions to switch to a "flexible" exchange rate regime and to liberalize trade and financial markets are expected to have had opposing effects on international trade and therefore potentially significant effects on import demand elasticities.[2]

The empirical model developed in this study has used the import demand equations of Watson and Teeucksingh (2000) and Modeste (2011) as a logical start point. Using annual data from 1970−1996 Watson and Teelecksingh (2000) modeled the log-level of real imports of non-durable of consumers' goods, a single equation of a small macroecenometric model, as a function of its lagged value and the contemporaneous value of real GDP. They found that consumers' imports were positively related to its lag and contemporaneous GDP, with coefficient estimates 0.56 and 0.84 respectively. Modest (2011) estimated the long- and short-run elasticities of the traditional and disaggregated expenditure models of aggregate goods imports using annual data from 1968−2006. The long-run relations in his analysis were determined using the bounds test for cointegration and the short-run elasticities determined in an error correction model. In the traditional model, he estimated imports of goods as a function of income and relative prices, where relative prices were computed as the ratio of the price of imports to the price of domestic goods. His long-run analysis revealed that income and relative prices were positively and negatively related to goods imports respectively. His short-run analysis found that contemporaneous changes in income and relative prices were positively and negatively related to changes in goods imports, with coefficient estimates 1.20 and -0.34 respectively. Long-run disequilibrium was estimated to have an adjustment coefficient of -0.21, suggesting roughly 20 percent was corrected each year.

The rest of this paper is organized as follows: Section 2 briefly describes the econometric methodology used to develop the VEqCM. Section 3 details the dataset and tests for integration and cointegration among the variables. Section 4 presents the dynamic system model and discusses the estimation results. Section 5 evaluates the dynamic system model and Section 6 concludes.

[1] Prior to April 13, 1993 when exchange rate controls were abolished and a flexible exchange rate regime was introduced, the CBTT maintained a fixed exchange rate with the US dollar. The exchange rate was devalued twice during the 1980s, the first was a 50 percent devaluation in December 1985 and the second was a lesser devaluation in August 1988. See Worrell *et al.* 2000 for a richer discussion on the exchange rate policy of Caribbean economies.

[2] Measuring the magnitudes of the opposing effects is in itself an interesting empirical exercise.

2 Econometric methodology

The econometric analysis commences with the modeling of the joint density of the stochastic variables. Hendry and Doornik (1994) discuss ten inter-related reasons concerning the logical and methodological basis for *Gets* modeling commencing from the joint density that include *inter alia* cointegration being a system property, being able to test weak, strong and super exogeneity and invariance, and being able to conduct multi-step ahead forecasts. The joint density of the vector of variables x_t is modeled as a vector autoregression (VAR) of the form

$$x_t = \sum_{j=1}^{s} \Pi_j x_{t-j} + \Phi q_t + v_t, \tag{1}$$

where $v_t \sim \mathrm{IN}[0,\Omega]$ and denotes an independent normal density with zero mean and covariance matrix Ω assumed to be (symmetric) positive definite, x_t is the vector of stochastic variables, q_t are deterministic terms, Π is the matrix of parameter estimates on lagged x_{t-j} and Φ is the vector of parameter estimates on q_t. The system can be represented in VEqCM form, which provides a useful reformulation when x_t are I(1) and retains the same basic innovation process v_t as the VAR, and is given by

$$\Delta x_t = \pi x_{t-1} + \sum_{j=1}^{s-1} \Pi_j \Delta x_{t-j} + \Phi q_t + v_t. \tag{2}$$

Letting $\pi = \alpha\beta'$, where α is a matrix of feedback coefficients and $\beta' x_{t-1}$ are I(0) cointegrating relationships, the VEqCM which serves as the start point for the econometric analysis is given by

$$\Delta x_t = \alpha\beta' x_{t-1} + \sum_{j=1}^{s-1} \Pi_j \Delta x_{t-j} + \Phi q_t + v_t. \tag{3}$$

3 Data properties

The empirical analysis uses seasonally unadjusted quarterly data spanning 1996(1)−2011(4). Five observations are lost due to differencing and lags resulting in an estimation sample spanning 1997(2)−2011(4). The CBTT and the Trinidad and Tobago Central Statistics Office (CSO), who jointly publish the Annual Statistical Digest and the Economic Bulletin, are the data sources for values of consumers' imports (M), liquidity, which is measured by the end of period M2 money supply (W) and output which is measured by the index of domestic production (Y). Note that liquidity is a stock variable measured as the end of period value, thus its one-period lag is used to obtain the start of period value ($W1$). The International Financial Statistics (IFS) published by the International Monetary Fund is the data source for both the TT and foreign (US) headline consumer price indices (CPI) (P) and (P^*) respectively. Values of consumers' imports and liquidity are in millions of nominal TT dollars per capita. The index of domestic production and both price indices are in base year 2005 $-$ the average of four quarters in 2005 = 100 $-$ and the ratio of relative prices is computed as ($Pr = P^*/P$). Each series is converted to its natural logarithm for use in estimation, with log levels denoted by lower-case letters.

Relative prices are an important series in the analysis of consumers' imports, and so deciding on appropriate measures for both the price index of consumers' imports and of domestic prices associated with non-durable consumers' goods is an important decision in the modeling exercise. Previous studies have used a measure of relative prices computed as the ratio of the imports price deflator to the index of domestic retail prices, albeit with annual data, see for example Watson (1990), Watson and Teelucksing (1997) and Maurin and Watson (2002) among others. Though it is possible to construct both a retail price index and producers' price index for a class of non-durable consumers' goods, this already arduous task is made only more so given data constraints.[3] Furthermore, although the TT retail price index and the US personal consumption expenditures index are available, these indices are not alike measures. Therefore, for the sake of consistency both the US and TT CPI are used in constructing the ratio of relative prices.

Several basic properties of the data series standout. Figure 1 displays graphs of the log-levels and log first-differences of the series. Visual inspection suggests the consumers' imports series displays more seasonality than the other series and examination of the autocorrelation functions (ACFs) and partial autocorrelation functions (PACFs) displayed in Figure 2 confirm this. Figure 1 also shows that consumers' imports grew at roughly the same rates as liquidity throughout the sample whereas it initially grew at a slower rate than output. Last, the rate of change in the relative price series increased mid-sample and is due almost entirely to the increased growth rate of the TT CPI.

3.1 Non-seasonal and seasonal integration

The preliminary univariate data analysis commences with an analysis of the seasonal properties of the data. Two test procedures are applied, namely that of Hylleberg *et al.* (1990) and Osborn *et al.* (1988). Hylleberg *et al.* (1990) (henceforth HEGY) propose a method to test for the presence

[3]Considering consumers' imports are traded at borders, the producer price index would be the preferred price index of domestic non-durable consumers' goods.

of seasonal and non-seasonal unit roots in quarterly time series based on the auxiliary regression

$$\phi(B)x_{4,t} = \mu_t + \pi_1 x_{1,t-1} + \pi_2 x_{2,t-1} + \pi_3 x_{3,t-2} + \pi_4 x_{3,t-1} + \epsilon_t, \tag{4}$$

where $\phi(B)$ is an AR polynomial of order $p - S$, μ_t is at most

$$\mu_t = \alpha_0 + \sum_{s=1}^{3} \alpha_s D_{st} + \beta_0 t,$$

and where

$$\begin{aligned} x_{4,t} &= (1 - B^4)x_t = \Delta_4 x_t, \\ x_{1,t} &= (1 + B + B^2 + B^3)x_t, \\ x_{2,t} &= -(1 - B)(1 + B^2)x_t, \\ x_{3,t} &= -(1 - B^2)x_t. \end{aligned}$$

The HEGY auxiliary regression can be estimated by ordinary lest squares (OLS), where the order of $\phi(B)$ is usually determined by diagnostic checks such that the estimated error process, $\hat{\epsilon}$, is approximately white noise. The appropriate differencing filter for x_t is determined by testing the significance of π_j, for $j = 1, 2, 3, 4$. When $\pi_2 = \pi_3 = \pi_4 \neq 0$ there are no seasonal roots present. When $\pi_1 = 0$ then the presence of the non-seasonal unit root 1 cannot be rejected.[4] When $\pi_2 = 0$ the presence of the seasonal unit root -1 cannot be rejected and when $\pi_3 = \pi_4 = 0$ the seasonal unit roots $\pm\, i$ cannot be rejected.

Table 1 reports the test results of applying the HEGY test procedure to the series m, y, $w1$, p, p^*, pr, $m - p^*$ and $w1 - p$. The test statistics are presented in the top panel and are evaluated using critical values tabulated in Frances and Hobijn (1997). The bottom panel reports the sample period, the number of lagged dependent variables and the impulse dummies used to whiten the errors in each test regression. Test results indicate there are no seasonal unit roots in any of the series; that the m, $w1$ and p, p^* pr, $m - p^*$ and $w1 - p$ series each have a non-seasonal unit root, i.e. are I(1,0); and that the y series may be trend stationary, i.e. may be I(0,0).

Osborn *et al.* (1988) (henceforth OCSB) and later Osborn (1990) propose a method to test for the presence two non-seasonal unit roots and seasonal unit roots. The primary distinction between this and the HEGY procedure is that the OCSB approach does not consider seasonal unit roots

[4] Note the HEGY test of $\pi_1 = 0$ is similar to the Augmented Dickey-Fuller test.

separately. It is based on the auxiliary regression

$$\phi(B)\Delta_1\Delta_S x_t = \mu_t + \pi_1(1 - B^S)x_{t-1} + \pi_2(1 - B)x_{t-S} + \epsilon_t, \tag{5}$$

where

$$\mu_t = \alpha_0 + \sum_{s=1}^{S-1} \alpha_s D_{st} + \beta_0 t + \sum_{s=1}^{S-1} \beta_s D_{ss} t.$$

The method first proposed by OCSB considers the case where all the β_0, ..., β_{S-1} terms are equal to zero. Frances and Koehler (1996) propose not setting these parameters equal to zero in order to select between models for time series with increasing seasonal variation. The OCSB auxiliary regression can be estimated by OLS, where similar to the HEGY procedure, the order of $\phi(B)$ is usually determined by diagnostic checks such that the estimated error process, $\hat{\epsilon}$, is approximately white noise. The appropriate differencing filter for x_t is determined by testing the significance of π_1 and π_2. When $\pi_1 = \pi_2 = 0$ the $\Delta_1\Delta_S$ filter is appropriate. When $\pi_1 = \pi_2 \neq 0$ no differencing filter is needed. When $\pi_1 = 0$ and $\pi_2 \neq 0$ the Δ_1, i.e. the (1 - B), filter is appropriate. And when $\pi_1 \neq 0$ and $\pi_2 = 0$ the Δ_S, i.e. the (1 - B^S), filter is appropriate.

Tables 2 reports the test results of the extended OCSB procedure for the case where β_1, ..., β_{S-1} are set equal to zero and Table 3 reports the test results for the case where no restrictions are imposed on the β coefficients. Similar to Table 1, the top panels in Tables 2 and 3 report the test statistics which are evaluated using critical values tabulated in Frances and Hobijn (1997), and the bottom panels report the sample period, the number of lagged dependent variables and the impulse dummies used to whiten the errors in each test regression. The test results in Tables 2 and 3 are similar apart for rejection of the null hypothesis of non-seasonal integration at the 5 percent level in the m and $m - p^*$ series when the auxiliary regression contains seasonal trend dummies. Overall, the results found by applying the OCSB procedure confirm the earlier results found using the HEGY procedure, i.e. there are no seasonal unit roots in any of the series and the m, $w1$, p, p^* pr, $m - p^*$ and $w1 - p$ series each have a non-seasonal unit root. Furthermore, the OCSB test results provide further evidence suggesting the output series, y, is trend stationary.

3.2 Cointegration

The multivariate data analysis commences with an investigation of the stationary cointegrating relationships of the system. The empirical analysis follows the VAR approach of Johansen (1988, 1991) and Johansen and Juselius (1990). It commences with a fifth-order VAR in the 4 stochastic variables ($m - p^*$, y, $w1 - p$, pr) with an intercept, trend, centered seasonal dummies and regime dummies, Dc, Dg and Ds.[5] The treatment of the deterministic variables, i.e. how they enter the system, is critical to a successful empirical analysis. In the following analysis only the trend term, t, was restricted to enter the cointegration space, leaving the intercept, centered seasonal dummies and regime dummies Dc, Dg and Ds unrestricted to be included in the short-run dynamics.[6]

Simplification tests on the initial VAR system suggested that a VAR(1) was sufficient for the present analysis, see Table 4 for reduction test statistics. The final VAR system thus uses 1 lag, an intercept, trend, centered seasonal dummies and regime dummies. The lag order was selected such that the system and single equation diagnostic tests indicated that the model specification was a satisfactory approximation to the unknown data generating process (DGP). In particular, the appropriateness of the specification was tested against the single equation and system variations of the Portmanteau test, the AR 1-4 test, The Jarque-Bera test for normality, the ARCH test for homoskedasticity and Ramsey's test for regression specification (RESET). Figures 3 and 4 display the system graphical diagnostics and recursive evaluation statistics respectively.

The determination of the number of cointegration vectors is based on the result of formal testing, the interpretability of the obtained coefficients of the eigenvectors, and graphical examination of the recursive eigenvalues. Table 5 reports the formal test results of the cointegration analysis, i.e. the eigenvalues and the associated trace (λ_{trace}) and maximum (λ_{max}) eigenvalue statistics along with the estimated eigenvectors and adjustment coefficient vectors, β and α respectively.[7] The result of formal testing supports the hypothesis that there are three cointegration vectors. The first eigenvector resembles an import demand relation with positive long-run effects from the deviation of output from trend and liquidity and a large negative effect from relative price levels.[8] The second eigenvector has most of its coefficient estimates near zero and appears to represent a trend stationary output relationship. The third partly resembles a liquidity (money) demand relation with a positive relation to the deviation of output from trend. The negative relation to relative prices, however, suggests a positive long-run effect from domestic prices which does not align with economic theories of money demand.[9] Interpretation of the coefficients of the eigenvectors therefore suggests there are two cointegration vectors.

[5] Dc is zero except for unity in 2008(4), 2010(2) and 2010(3) and adjusts for the largest residual outliers attributed to the recent financial crises. Dg is zero except for $\pm$ unity in 2003(4) and 2004(1) respectively and is attributed the the commencement of large scale production of liquified natural gas in TT which led to the economy being re-classified as a gas-based economy as opposed to an oil-based economy. Ds is zero except for unity in 2011(3) and 2011(4) and is attributed to the policy imposed curfew that lasted from August 21st, 2011 to December 5th, 2011.

[6] Setting the trend unrestricted yields similar cointegration results but the interpretation is not as straightforward.

[7] The multivariate stationarity test statistics indicate stationarity is rejected for all series save for output but only when the coefficient on trend is allowed to enter the long-run relation unrestricted. These results reconfirm the integration test results reported in Section 3.1.

[8] Several authors have found similar results regarding coefficient signs in studies examining imports, output and relative prices among other variables, see for example Watson and Teelucksing (1997) and Modeste (2011).

[9] Note the trend stationary relation of the second eigenvector is also present in the first and third eigenvectors.

Figure 5 plots the recursive eigenvalues for the estimated long-run relations. The first recursive eigenvalue is decreasing till mid-sample, but becomes constant and is clearly non-zero; the second is also constant from mid-sample and is also clearly non-zero; the third, however, is non-constant throughout the sample, near zero in the mid-sample and only increases at the end of the sample; and the fourth is everywhere close to zero. Graphical examination of the recursive eigenvalues supports the hypothesis that there are two *constant* cointegration vectors.

The test statistics, interpretation of the eigenvectors and graphical examination support hypothesis that there are two cointegration vectors. Therefore, subsequent analysis, results and discussions will be based on the assumption of two restricted cointegration vectors which are determined by excluding the regime dummies, intercept and centered seasonal dummies, but including the trend.

The following identification restrictions, reported in Table 6, were imposed on the unrestricted cointegration vectors: for long-run consumers' imports the trend stationary output relation of the second cointegration vector and an opposing effect from liquidity one-fourth that of relative prices estimated by the model, i.e. for an order of $\beta' = (m - p^*, y, w1 - p, pr, t)$ restrictions were $\beta' = (1, -a, -b, 4b, ac)$; and for output zero restrictions on all stochastic variables leaving the trend coefficient to be estimated by the model, i.e. $\beta' = (0, 1, 0, 0, -c)$. The estimated values for a, b and c were 0.874, 0.299 and 0.023 respectively. Figure 6 displays the data mapped to the I(0) linear combinations defining the error correction terms given by:

$$ecm1_t = (m - p^*)_t - (0.874y_t - 0.020t) - 0.299(w1 - p)_t + 1.197pr_t \tag{6}$$

and

$$ecm2_t = y_t - 0.023t, \tag{7}$$

It is worth noting that the first cointegration vector as posited by hypothesis H_1 appears to be the linear combination of two stationary processes, namely a stationary relation among consumers' imports, liquidity and relative prices and the trend stationary output relation.[10] Formal test results reported in Table 6 support this claim. Specifically, hypothesis H_{a1} for the consumers' imports cointegration vector wherein $a = 0$, i.e. the coefficient on output and by construction also the coefficient on trend, is not rejected; and hypothesis H_{a2} wherein the coefficients on the output and trend variables enter the cointegration relation unrestricted yields a long-run relation identical to that found when the trend stationary output relation is imposed. Interpreted individually, the first cointegration vector proposed by hypothesis H_{a1} suggests a positive long-run effect from liquidity

[10]Note that any linear combination of the stationary vectors is also a stationary vector and therefor direct interpretation of individual cointegration vectors is not always interesting, see for example Johansen and Juselius (1992).

and a large negative effect from relative price levels but no long-run effect from output. This relation does not appear very interesting on its own seeing as output has no long-run effect. Considering both cointegration vectors jointly, however, suggests a more meaningful economic relation wherein both error/proportional and integral control mechanisms influence economic behavior. Several empirical studies on consumers' expenditure found evidence of both error and integral control mechanisms, see *inter alia* Hendry and von Ungern-Sternbeg (1981) and Campos and Ericsson (1988); and it is reasonable to conjecture that both mechanisms would play important roles in cases where consumers' imports constitute a large share of consumer's expenditure.[11] Since the cointegration relation which includes the trend stationary relation appears the more economically meaningful stationary relation subsequent analysis will be based on this as the first stationary relation.

The restricted cointegration vectors were tested for lying in the cointegration space jointly with testing for long-run weak exogeneity of $(y, w1-p, pr)$ for the consumers' imports function, i.e. the long-run consumers' imports relation not appearing in any of the other stochastic equations; and $(m-p^*, w1-p, pr)$ for the long-run trend stationary output relation, which would suggest the system could be reduced to a 3-dimensional system by conditioning on y if this hypothesis is not rejected. The long-run weak exogeneity tests involved testing restrictions on the adjustment coefficient vectors, $\alpha's$, with the eigenvectors, β', identified as in (6)$-$(7). For example, testing the long-run weak exogeneity of the consumers' imports function on the remaining stochastic variables had the form $\alpha' = (*, 0, 0, 0)'$. Formal test results presented in Table 7 suggest the trend stationary output relation is weakly exogenous for all other stochastic variables; and the consumers' imports relation is long-run weakly exogenous for output and liquidity but not relative price variables. The result suggesting the consumers' imports relation is not long-run weakly exogenous for relative prices is in line with the findings of Primus et al. (2011) and Mahabir and Jagessar (2011) wherein foreign prices Granger-cause domestic prices and there is imported inflation in TT prices. Altogether these results suggest a conditional (single equation) model of consumers' imports will not suffice, but rather modeling the linear dynamic system would be appropriate, see Hendry and Doornik (1994).

4 General to specific modeling of the system

To establish a baseline innovation variance the system modeling exercise commences with estimating the 6-equation system in $(d(m-p^*)_t, dy_t, d(w1-p)_t, dpr_t, ecm1_t, ecm2_t)$.[12] Here, the error correction

[11] Integral control mechanisms (ICMs) are most easily interpreted by introduction of a state variable defined by using end-of-period definitions, such as liquid balances. In terms of the other variables in the system, the ICM is the integral of past discrepancies between the choice variables in the error correction mechanism, for example consumption and income/output.

[12] Given the long-run weak exogeneity of y and $w1-p$ for the parameters of the import demand equation it would suffice to condition on dy and $d(w1-p)$ in the dynamic system. The system is preserved, however, as this readily

terms are defined by identities as such a representation allows for a more interpretable model. The identities for the error correction terms follow from their construction in (6)−(7) and are given by:

$$ecm1_t \equiv ecm1_{t-1} + d(m-p^*)_t - (0.874dy_t - 0.020dt) - 0.299d(w1-p)_t + 1.197dpr_t \tag{8}$$

and

$$ecm2_t \equiv ecm2_{t-1} + dy_t - 0.023dt, \tag{9}$$

The available information set for system estimation therefore comprises $((d(m-p^*)_{t-i}, dy_{t-i}, d(w1-p)_{t-i}, dp^*_{t-i}, dp_{t-i}, ecm1_t, ecm2_t))$ for $i = 1, ..., 4$, together with an intercept, centered seasonal dummies and regime dummies Dc, Dg and Ds. Recall the trend was restricted to enter the cointegration space and so is not included in the current information set. During the short-run analysis the relative price variables dpr_{t-i} were separated into dp^*_{t-i} and dp_{t-i} to test for asymmetric effects of domestic and foreign inflation and also as there was no *a priori* reason to impose these restrictions in the short-run. Tables 8−11 display the system regression estimates and Table 12 the residual correlations of the of the general unrestricted model (GUM).

Reduction from the GUM to the final parsimonious model occurs in two steps. First the *Autometrics* algorithm of *PcGive* is applied equation by equation.[13] Second, nonzero parametric restrictions were imposed to arrive at a more parsimonious and economically interpretable specification. In particular, the restriction of equality on the four lags of output growth in the consumers' imports equation, which may be interpreted a statistical smoothing of output in order to extract changes that are more permanent; and of equality on the first and third lags of foreign inflation in the relative price equation that implies the term $(d^2p^*_{t-1} + d_2p^*_{t-2})$, which may be interpreted as both a statistical smoothing of past and a data-based predictor of future foreign inflation. The reductions result in the following system specification:

$$\begin{aligned}
\text{d(m-p*)} \;=\; & -\underset{(0.099)}{0.24}\ \text{d(m-p*)}_{t-1} + \underset{(0.21)}{0.7}\ \text{d}_4\text{y}_{t-1} - \underset{(0.24)}{1.1}\ \text{d(w1-p)}_{t-3} \\
& + \underset{(1.6)}{4.7}\ \text{dp*}_{t-2} + \underset{(0.73)}{1.6}\ \text{dp}_{t-4} - \underset{(0.093)}{0.24}\ \text{ecm1}_{t-1} \\
& - \underset{(1.2)}{3.1} - \underset{(0.031)}{0.31}\ \text{CS}_t - \underset{(0.037)}{0.056}\ \text{CS}_{t-1} - \underset{(0.029)}{0.045}\ \text{CS}_{t-2}
\end{aligned}$$

allows for testing strong and super exogeneity.

[13] The following *Autometrics* settings were used: target size was set to 0.10; no outlier and break detection; pre-search lag reduction and pre-search variable reduction were turned on; search effort was set to 1; backtesting was set to GUM 0; the tie breaker was the Schwarz Criterion; diagnostic test *p-value* was set to 0.01; and GIVE was set to do the reduced form first. Note the fourth lag of the change in liquidity, i.e. $d(w1-p)_{t-4}$ was set unrestricted in the liquidity equation, and lead to a more sensible final equation.

$$
\begin{aligned}
\text{dy} \;=\; & -\underset{(0.098)}{0.25}\ \text{dy}_{t-3} - \underset{(0.12)}{0.69}\ \text{ecm2}_{t-1} - \underset{(0.022)}{0.14}\ \text{Ds}_t \\
& + \underset{(0.45)}{2.5} - \underset{(0.013)}{0.013}\ \text{CS}_t + \underset{(0.012)}{0.016}\ \text{CS}_{t-1} + \underset{(0.012)}{0.014}\ \text{CS}_{t-2}
\end{aligned}
$$

$$
\begin{aligned}
\text{d(w1-p)} \;=\; & \underset{(0.035)}{0.069}\ \text{d(m-p*)}_{t-2} - \underset{(0.1)}{0.25}\ \text{d(w1-p)}_{t-4} - \underset{(0.65)}{1.1}\ \text{dp*}_{t-1} \\
& - \underset{(0.017)}{0.043}\ \text{Dc}_t + \underset{(0.021)}{0.1}\ \text{Dg}_t + \underset{(0.0058)}{0.026} \\
& + \underset{(0.012)}{0.031}\ \text{CS}_t - \underset{(0.011)}{0.02}\ \text{CS}_{t-1} + \underset{(0.016)}{0.0073}\ \text{CS}_{t-2}
\end{aligned}
$$

$$
\begin{aligned}
\text{dpr} \;=\; & \underset{(0.012)}{0.04}\ \text{d(m-p*)}_{t-1} - \underset{(0.027)}{0.087}\ \text{dy}_{t-2} - \underset{(0.15)}{0.6}\ (\text{d}^2\text{p*}_{t-1}+\text{d}_2\text{p*}_{t-2}) \\
& - \underset{(0.011)}{0.044}\ \text{ecm1}_{t-1} - \underset{(0.0047)}{0.049}\ \text{Dc}_t - \underset{(0.13)}{0.55} \\
& + \underset{(0.0032)}{0.0047}\ \text{CS}_t + \underset{(0.0039)}{0.013}\ \text{CS}_{t-1} - \underset{(0.0037)}{0.012}\ \text{CS}_{t-2}
\end{aligned}
\tag{10}
$$

log-likelihood = 542.82 -T/2log|Omega| = 877.69
no. of observations = 59 no. of parameters = 35
LR test of over-identifying restrictions: $\chi^2(93) = 84.21$ [0.731]

System summary statistics are reported under (10), Table 13 reports the correlation of structural residuals, Table 14 reports system diagnostics using the reduced-form residuals and Figure 7 displays the graphical system diagnostics. The summary statistics clearly indicate a good fit of the system, with a system Likelihood Ratio (LR) test of over-identifying restrictions statistic of 84.21 (*p-value* = 0.731) and individual regression standard errors of 0.071, 0.030, 0.029, 0.008 for equations $d(m-p^*)$, dy, $d(w1-p)$ and dpr respectively. None of the diagnostic tests are rejected for the individual equations or the system save for the heteroskedasticity test for the liquidity equation which is rejected at the 5 percent significance level. Altogether, these results indicate that the VEqCM given by (10) is statistically well specified. Moreover, the recursive estimates portrayed in Figure 8 depict empirically constant parameters and provide evidence of the model's robustness.

Estimation results for the change in consumers' imports equation indicate significant short-run effects from consumers' imports, output and liquidity growth, foreign and domestic inflation, and moderate adjustment to long-run disequilibrium. Specifically, changes in consumers' imports are negatively correlated with an estimated coefficient of -0.24; positively related to *smoothed* output growth and both foreign and domestic inflation, albeit at different lags, with estimated coefficients 0.7, 4.7 and 1.6 respectively; and negatively related to liquidity growth with an estimated coefficient of -1.1. The estimated short-run elasticities for foreign and domestic inflation suggests that consumers are more concerned with the former and also that it affects their decision making and/or enters their information set sooner than the latter. The estimated long-run adjustment coefficient suggests that approximately 25 percent of long-run disequilibrium is adjusted each quarter. This estimate is notably larger than those found in previous studies which are generally near 20 percent per year, though only for a subsection of total goods imports, see for example Modeste (2011).

Estimation results for the remaining system equations also provide some explanatory power but it is the relative price series that appears most interesting among them in the current analysis. Output growth is negatively autocorrelated, responds strongly to deviations from its long-run trend and as expected was influenced by the policy imposed curfew of the last two quarters of 2011, Ds. Liquidity growth depends on lagged changes in consumers' imports, liquidity and foreign inflation, and was influenced by regime dummies associated with the financial crises, Dc, and the onset of large scale production of liquified natural gas in TT, Dg. Changes in relative prices depends on the term $(d^2p^*_{t-1} + d_2p^*_{t-2})$, somewhat on lagged changes in consumers' imports and output and very little to long-run disequilibrium in consumers' imports. Note that domestic inflation is not an explanatory variable for relative prices, which may seem surprising at first given that foreign inflation is essentially an exogenous variable in the current analysis. Inspection of the CBTT monetary policy decisions, however, indicate that expectations of US monetary policy is one of the key factors taken into consideration when setting TT monetary policy and thus lends some credence to this result.[14]

5 Evaluating the model

5.1 Parameter constancy

Figure 8 reports for the latter part of the sample the recursive FIML coefficient estimates along with the respective 1-step residuals, ± 2 standard error bands, and the 1-step ahead ($1up$) and Break-point Chow statistics. Recursive coefficient estimates show little variation throughout the available estimation period, and $1up$ and Break-point Chow statistics are all insignificant. These statistics all point to the empirical constancy of the model's parameters.

[14]See CBTT Monetary Policy Decisions.

5.2 Exogeneity

The strong and super exogeneity of output and liquidity were tested for the parameters of the consumers' import equation. Specifically, the significance of parameter estimates on lagged output and liquidity growth in the consumers' import equation were evaluated to test the Granger non-causality of these variables. Estimation results reported in (10) indicate statistically significant feedback from these variables and hence strong exogeneity is rejected in both cases.

The *significant* impulse dummies found in the equations for output and liquidity growth were tested for relevance in the change in consumers' import equation. Estimation results reported in (10) point to the irrelevance of the regime dummies significant to both output growth, Ds, and liquidity growth, Dc and Dg, and thus super exogeneity is not rejected for either variable.

Together, the strong and super exogeneity test results suggest that multistep-ahead forecasting of the consumers' import equation is not recommended outside of a system specification and that valid inference from counterfactual simulations may be drawn from those altering future paths of both output and liquidity.

5.3 Encompassing

The model developed in this study was evaluated to determine its ability to account for the results of the models of Watson and Teelucksingh (2000) and Modeste (2011). Watson and Teelucksingh (2000) modeled the log of consumers' imports as a function of its lagged value, contemporaneous and lagged values of real GDP and the ratio of the import price index relative to the retail price index using annual time series data for 1970−1996. Modest (2011) estimated elasticities for the traditional and disaggregated expenditure models of import demand for three caribbean countries, including TT, using the bounds test for cointegration and annual time series data for TT for 1968−2006.[15] The unrestricted models estimated by Watson and Teelucksingh (2000) and Modeste (2011), for the traditional model, are given respectively by:

$$(m-p^*)_t = \alpha_0 + \sum_{i=1}^{k} \alpha_{1i}(m-p^*)_{t-i} + \sum_{i=0}^{k} \alpha_{21} y_{t-i} + \sum_{i=0}^{k} \alpha_{3i} pr_{t-i} + \varepsilon_t, \tag{11}$$

[15] The traditional model of import demand models imports as a function of income and relative prices, whereas the disaggregated expenditure model models imports as a function of private and public consumption expenditure, expenditure on investment goods, and receipts from exports.

$$d(m-p^*)_t = \beta_0 + \sum_{i=1}^{k} \beta_{1i} d(m-p^*)_{t-i} + \sum_{i=0}^{k} \beta_{21} dy_{t-i} + \sum_{i=0}^{k} \beta_{3i} dpr_{t-i} + \beta_4 ecm_{M1,t-1} + \varepsilon_t. \quad (12)$$

In the above equations, $m-p^*$ represents real consumers' imports in (11) and real imports in (12), y represents real GDP in (11) and real income in (12), pr represents relative prices and ecm_{M1} is an error correction term.[16] Estimation results reported in the respective articles were[17]

$$\text{m-p*}_t = \underset{(2.943)}{-\ 5.276} + \underset{(0.146)}{0.562}\ \text{m-p*}_{t-1} + \underset{(0.359)}{0.838}\ \text{y}_t,$$

$\hat{\sigma} = 0.1513$ $\quad R^2 = 0.749$ $\quad$ Adj. $R^2 = 0.724$ $\quad$ RSS $= 0.458$ $\quad$ F-statistic $= 29.807$

and

$$\text{d(m-p*)}_t = \underset{(0.023)}{-\ 0.018} + \underset{(0.326)}{1.203}\ \text{dy}_t - \underset{(0.161)}{0.343}\ \text{dpr}_t - \underset{(0.066)}{0.207}\ \text{ecm}_{M1,t-1},$$

Adj. $R^2 = 0.97$ $\quad$ F-statistic $= 9.861$ $\quad$ AR $= 4.024$ [0.258] $\quad$ Normality $= 0.355$ [0.836]

where the long-run elasticities for the error correction term ecm_{M1} were estimated as

$$\text{m-p*}_t = \underset{(0.908)}{2.383} + \underset{(0.356)}{1.255}\ \text{y}_t - \underset{(0.09)}{0.384}\ \text{pr}_t.$$

The models of Watson and Teelucksingh (2000), Modeste (2011) and the specification of the consumers' imports equation developed in this paper can be viewed as contending theories nested in the general *single equation* model given by

$$d(m-p^*)_t = \sum_{i=1}^{k} \eta_{mi} d(m-p^*)_{t-i} + \sum_{z} \sum_{i=0}^{k} \eta_{zi} dz'_{t-i} + \alpha\beta'(m-p^*, z')_{t-1} + \Phi q_t + \varepsilon_t, \quad (13)$$

where $z' = (y, w1-p, pr)$, η_{ji} for $j = m, z$ are parameter estimates on the stochastic terms, q_t are deterministic terms and Φ is the vector of parameter estimates on q_t. Using (15) as a start point, it is now possible to perform encompassing and misspecification tests.

Equations (11) and (12) were re-estimated using the current quarterly dataset, including centered

[16] Watson and Teelucksingh (2000) compute real consumers' imports as the nominal TT dollar value deflated by the imports of non-durable consumer goods deflator.

[17] Note the diagnostic statistics reported under each equation were the statistics made available in each publication.

seasonal dummies, and produced the following respective results:[18]

$$\begin{aligned} \text{d(m-p*)} \;=\; & -\underset{(1.4)}{5.6} - \underset{(0.096)}{0.37}\ \text{m-p*}_{t-1} + \underset{(0.07)}{0.27}\ \text{y}_t \\ & - \underset{(0.034)}{0.29}\ \text{CS}_t - \underset{(0.036)}{0.029}\ \text{CS}_{t-1} - \underset{(0.033)}{0.065}\ \text{CS}_{t-2}, \end{aligned}$$

T = 59 $\hat{\sigma} = 0.090$ $R^2 = 0.747$ Adj. $R^2 = 0.724$ RSS = 0.426
F(5,53) = 31.37** [0.000] AR: F(4,49) = 1.000 [0.417] ARCH: F(4,51) = 1.227 [0.311]
Normality: $\chi^2(2) = 0.790$ [0.674] Hetero: F(7,51) = 1.663 [0.139]
Hetero-X: F(14,44) = 1.333 [0.228] RESET23: F(2,51) = 0.894 [0.415]

and

$$\begin{aligned} \text{d(m-p*)} \;=\; & -\underset{(1.5)}{6.2} - \underset{(0.29)}{0.048}\ \text{dy}_t - \underset{(0.84)}{0.16}\ \text{dpr}_t - \underset{(0.1)}{0.42}\ \text{ecm}_{M1,t-1} \\ & - \underset{(0.036)}{0.29}\ \text{CS}_t - \underset{(0.036)}{0.023}\ \text{CS}_{t-1} - \underset{(0.033)}{0.066}\ \text{CS}_{t-2}, \end{aligned}$$

T = 59 $\hat{\sigma} = 0.089$ $R^2 = 0.754$ Adj. $R^2 = 0.726$ RSS = 0.414
F(6,52) = 26.59** [0.00] AR: F(4,48) = 0.875 [0.486] ARCH: F(4,51) = 1.216 [0.316]
Normality: $\chi^2(2) = 0.865$ [0.649] Hetero: F(9,49) = 1.305 [0.259]
Hetero-X: F(21,37) = 0.854 [0.643] RESET23: F(2,50) = 0.888 [0.418]

where re-estimation of long-run elasticities for ecm_{M1} yielded[19,20]

$$\text{m-p*} = -\underset{(1.019)}{14.836} + \underset{(0.231)}{0.653}\ \text{y}_t - \underset{(0.557)}{0.13}\ \text{pr}_t - \underset{(0.236)}{0.709}\ \text{CS}_t - \underset{(0.086)}{0.055}\ \text{CS}_{t-1} - \underset{(0.087)}{0.159}\ \text{CS}_{t-2}.$$

The consumers' imports equation of the VEqCM was re-estimated as a single equation by OLS

[18] Note the estimated elasticities on lagged imports in Watson and Teelucksing (2000) and contemporaneous output growth in Modeste (2011) have opposite signs to those found using the annual data samples in the respective studies.

[19] The long-run solutions were solved using a first-order ADL. Initially a fifth-order ADL that included an intercept and centered seasonal dummies was estimated but reduction to the first-order ADL was preferred by the SC, HQ and AIC.

[20] Several ADL specifications were estimated to derive static long-run solutions used to create the ecm_{M1} term, for example without an intercept and/or centered seasonal dummies. The specification reported contains an unrestricted intercept and unrestricted centered seasonal dummies. All specifications led to similar encompassing test results.

and produced the following parameter estimates

$$
\begin{aligned}
\text{d(m-p*)} \quad = \quad & -\underset{(0.1)}{0.24}\ \text{d(m-p*)}_{t-1} + \underset{(0.23)}{0.75}\ \text{d}_4\text{y}_{t-1} - \underset{(0.26)}{1}\ \text{d(w1-p)}_{t-3} \\
& + \underset{(1.6)}{4.7}\ \text{dp*}_{t-2} + \underset{(0.77)}{1.8}\ \text{dp}_{t-4} - \underset{(0.097)}{0.22}\ \text{ecm1}_{t-1} \\
& - \underset{(0.031)}{0.31} - \underset{(1.2)}{2.8}\ \text{CS}_t - \underset{(0.039)}{0.05}\ \text{CS}_{t-1} - \underset{(0.029)}{0.046}\ \text{CS}_{t-2}
\end{aligned}
$$

T = 59 $\hat{\sigma} = 0.072$ $R^2 = 0.850$ Adj. $R^2 = 0.823$ RSS = 0.252
F(9,49) = 30.91** [0.000] AR: F(4,45) = 0.247 [0.910] ARCH: F(4,51) = 1.034 [0.399]
Normality: $\chi^2(2) = 0.562$ [0.755] Hetero: F(15,43) = 1.165 [0.334]

Table 15 reports the encompassing and misspecification test results for the models of Watson and Teelucksing (2000) and Modeste (2011) against the specification developed in this paper, labeled WT, M and B respectively. In particular, the tests of Cox (1961), Ericsson (1983), Sargan (1980) and a joint model *F-test* are used to test the encompassing results. Results suggest the specification developed in this study encompasses the models of Watson and Teelucksing (2000) and Modeste (2011) for the specified sample period.[21] The specification developed in this study thus makes an important contribution in the progressive research towards understanding consumers' imports.

5.4 Forecasting

The VEqCM of consumers' imports developed in this paper demonstrates desirable statistical properties and appears to be a sufficient approximation of the underlying DGP. These reasons make the model a potentially useful tool for ex ante and ex post forecasts, and policy simulations at least with respect to output and liquidity. 1-step ahead forecasting from this model is straightforward given the current quarter change in the stochastic variables depends only on lagged information and multistep-ahead forecasting is also possible given the system specification. Ex posts forecasts which are often used to asses the constancy of a model, see for example de Brouwer and Ericsson (1998) and Hendry and Doornik (1994), are however only considered for 1-step ahead forecasts.

[21] The model developed in this study parameter encompasses the models of Watson and Teelucksing (2000) and Modeste (2011). Given the quarterly dataset, versions of the latter models with enriched dynamics selected by *Autometrics* were also estimated to test encompassing results with test results unchanged. Furthermore, abstracting from the nonzero parametric restrictions applied to the model developed in this paper led to more favorable encompassing test results.

For the ex post forecasting exercise (10) is re-estimated over the subsample 1997(2)−2009(4) and the re-estimated parameters are then used to produce forecasts over the period 2010(1)−2011(4).[22] Figure 9 displays the actual, fitted and forecast values of changes in consumers' imports, output, liquidity and relative prices, with ± 2 standard error bars for each forecast. The forecasts track the realized values of consumers' imports well with all point forecasts falling within the standard error bars and near the realized values. The accuracy of these forecasts are in spite of both the financial crisis and policy imposed curfew regimes, and thus further attests to the constancy of the model for consumers' imports developed in this paper. Moreover, the model successfully predicts the directional change of the point estimate in all but one forecast, specifically 2010(4) which is the first point forecast after the end of the financial crisis regime. This result suggests the model is well suited for forecasting consumers' imports but that caution must be made when exiting and possibly entering a regime that would affect the demand for liquidity and/or inflation.

The VEqCM also forecasts the remaining stochastic variables fairly accurately, with the caveats that it does not anticipate the policy regime that affected output and that which initially signaled the *end* of the financial crisis. As a result, forecasts for output fall outside the respective 95% confidence intervals for the final two observations of the forecast period, i.e 2011(3) and 2011(4) whereas they otherwise tracked the series well; and forecasts for relative prices fall outside the confidence interval for the first period indicating the *end* of the financial crisis, i.e 2010(2). Notably, the liquidity series does not encounter any of the issues that hinder the performance of the other series as all forecasts fall within the confidence intervals though the forecast for 2009(1) sits virtually at the confidence interval. Whether or not this last result is spurious, however, is untested.

6 Concluding remarks

This this paper examines liquidity effects on the demand for imports of non-durable consumers goods. The case of Trinidad and Tobago is examined as it fits the description of both a small island developing state and a small non-renewable natural resource economy, both small open economies for which consumers' expenditure on imports are likely to represent a large share of aggregate consumers' expenditure. A parsimonious vector equilibrium correction model (VEqCM) is used to test the hypotheses that liquidity has both long- and short-run effects on consumers imports, and whether there are asymmetric foreign and domestic price effects.

Liquidity, which is measured as the start of period M2 money supply, is found to have both long- and short-run effects on imports of non-durable consumers' goods. Cointegration analysis reveals

[22]Note, no location shifts are used in the forecasting exercise.

a long-run relation among consumers' imports, the deviation of output from trend, liquidity and the ratio of foreign relative to domestic prices. Specifically, output and liquidity are found to be positively related to consumers' imports with elasticities near 0.9 and 0.3 respectively; and relative prices negatively related with an elasticity near -1.2. Note that despite employing a measure of income different from that commonly used in the literature, i.e. the domestic production index as opposed to real GDP or real income, the long-run effects of income are still found to be significant. In the short-run liquidity is found to have a lagged negative effect on consumers' imports.

General to specific modeling reveals negative autocorrelation from consumers' imports, asymmetric foreign relative to domestic price effects, a significant role for income, and moderate disequilibrium adjustment. Results indicate both foreign and domestic prices have lagged positive effects on consumers' imports, albeit at different lags. The previous four quarters of output growth were found to be positively related to changes in consumers' imports in a manner that may suggest consumers are more concerned with output changes that are more permanent. Long-run disequilibrium adjustment was estimated at approximately 25 percent each quarter. This estimate is notably larger than those found in previous studies, which were generally near 20 percent per year, albeit for a long-run relation that abstracted from liquidity and that focused on total goods imports.

The VEqCM developed in this study improves on previous specifications, demonstrates desirable statistical properties and appears a sufficient approximation of the underlying DGP. Recursive estimates and 1-step ahead ex-post forecasts demonstrate the empirical constancy of the system's parameters. Strong exogeneity test results suggest valid inference on multi-step forecasting is only possible in a system specification and super exogeneity test results suggest valid inference on policy simulations may only be made regarding output and liquidity. Encompassing test results indicate the specification developed in this paper parameter encompasses previous models and highlights the important role of liquidity, and indirectly monetary policy, in understanding consumers' imports.

7 References

Campos, J. and N.R. Ericsson, 1988, Econometric modeling of consumers expenditure in Venezuela, International Finance Discussion Paper 325 (Board of Governors of the Federal Reserve System, Washington, DC).

Campos, J., N.R. Ericsson and D.F. Hendry, 1996, Cointegration tests in the presence of structural breaks, Journal of Econometrics 70, 187−220.

Chow, G.C., 1960, Tests of equality between sets of coefficients in two linear regressions, Econometrica 28, 591−605.

Cox, D. R., 1961, Tests of separate families of hypotheses, in: Proceedings of the Fourth Berkeley Symposium on Mathematical Statistics and Probability, Volume 1, Berkeley, 105−123 (University of California Press, California).

Davidson, J.E.H, D.F. Hendry, F. Srba and S. Yeo, 1978, Econometric modelling of the aggregate time-series relationship between consumers' expenditure and income in the United Kingdom, Ch. 8 in: D.F. Henry, ed., Econometrics: Alchemy or science (Oxford University Press, Oxford).

Doornik, J.A. and D.F. Hendry, 2013, PcGive 14.0 (Timberlake Consultants Ltd., London).

Engle, R. and C.W.J. Granger, 1987, Cointegration and error correction: Representation, estimation and testing, Econometrica 55, 251−76.

Ericsson, N.R., 1983, Asymptotic properties of instrumental variables statistics for testing non-nested hypotheses, Review of Economic Studies 50, 287−303.

Ericsson, N.R., 1992, Cointegration, exogeneity, and policy analysis: An overview, Journal of Policy Modeling 14, 251−280.

Ericsson, N.R., 1995, Conditional and structural error correction models, Journal of Econometrics 69, 159−171.

Ericsson, N.R., and J.S. Irons, eds., 1994, Testing exogeneity (Oxford University Press, Oxford).

Frances, P.H., 1992, Testing for seasonality, Economic Letters 38, 259−262.

Franses, P.H. and B. Hobijn, 1997, Critical values for unit root tests in seasonal time series, Journal of Applied Statistics 24, 25−47.

Franses, P.H. and A.B. Koehler, 1996, Model selection strategies for time series with increasing seasonal variation, (revised version of) Econometric Institute Report 9308 (Erasmus University Rotterdam, Rotterdam).

Hendry, D.F. and J.A. Doornik, 1994, Modelling linear dynamic econometric systems, Scottish Journal of Political Economy 41, 1−33.

Hendry, D.F. and T. von Ungern-Sternberg, 1981, Liquidity and inflation effects on consumers' expenditure, Ch. 9 in: D.F. Hendry, ed., Econometrics: Alchemy or science (Oxford University

Press, Oxford).

Hilaire, D.L., S.M.A Nicholls, and A.J. Henry, 1990, Forecasting and policy evaluation in a macroeconometric model of the Trinidad and Tobago economy, Social and Economic Studies 39, 107−148.

Holder, C and O. Williams, 1997, Trinidad and Tobago: Managing import demand during oil shocks, in: D.Worrell and R. Craigwell, eds., Macroeconomics and finance in the Caribbean: Quantitative analyses (Caribbean Centre for Monetary Studies, University of the West Indies) 55−81.

International Monetary Fund, Statistics Department, 2014, International financial statistics (International Monetary Fund, Washington, DC).

Johansen, S., 1988, Statistical analysis of cointegration vectors, Journal of Economic Dynamics and Control 12, 231−254.

Johansen, S., 1992, Testing weak exogeneity and the order of cointegration in UK money demand, Ch. 5 in: N.R. Ericsson and J.S. Irons, eds., Testing exogeneity (Oxford University Press, Oxford).

Johansen, S., 1991, Estimation and hypothesis testing of cointegration vectors in Gaussian vector autoregressive models, Econometrica 59, 1551−1580.

Johansen, S. and K. Juselius, 1990, Maximum likelihood estimation and inference on cointegration − With applications to the demand for money, Oxford Bulletin of Econometrics and Statistics 52, 169−210.

Johansen, S. and K. Juselius, 1992, Testing structural hypotheses in a multivariate cointegration analysis of the PPP and UIP for UK, Journal of Econometrics 53, 211−244.

Juselius, K., 1992, Domestic and foreign effects on prices in an open economy: The case of Denmark, Ch. 7 in: N.R. Ericsson and J.S. Irons, eds., Testing exogeneity (Oxford University Press, Oxford).

Mahabir, R. and V. Jagessar, 2011, An Examination of the Import Price Transmission Mechanism in Trinidad and Tobago, Central Bank Working Paper 02/2011 (The Central Bank of Trinidad and Tobago, Port of Spain).

Maurin, A. and P.K. Watson, 2002, Quantitative modelling of the Caribbean macroeconomy for forecasting and policy analysis: Problems and solutions, Social and Economic Studies 51, 1−47.

Nelson C. Modeste, 2011, An Empirical Analysis of the Demand for Imports in Three CARICOM Member Countries: An Application of the Bounds Test for Cointegration, Review of Black Political Economy, 38, 53−62.

Osborn, D.R., 1990, A survey of seasonality in UK macroeconomic variables, International Journal of Forecasting 6, 327−336.

Osborn, D.R., A.P.L Chui, J.P. Smith and C.R. Birchenhall, 1988, Seasonality and the order of integration for consumption, Oxford Bulletin of Economics and Statistics 50, 361−377.

Primus, K., V. Jagessar, D. Cox, and R. Mahabir, 2011, What Accounts for Food Price Inflation in Trinidad and Tobago in Recent Years, Economic Bulletin, XIII, no 1, 89−98 (The Central Bank of Trinidad and Tobago, Port of Spain).

Sargan, J. D., 1980, Some tests of dynamic specification for a single equation, Econometrica 48, 879−897.

Trinidad and Tobago, Central Statistical Office, 2015 (Port of Spain, Trinidad, WI).

Trinidad and Tobago, Central Bank of Trinidad and Tobago, 2015 (Port of Spain, Trinidad, WI).

Watson, P.K., 1990, Modelling the balance of payments of Trinidad and Tobago 1965−1985: A policy making perspective, Social and Economic Studies 39, 51−70.

Watson, P.K., 1997, Forecasting macroeconomic time series in the Caribbean: VAR and univariate ARIMA models compared, Paper presented at the 17th Annual International Symposium on Forecasting, Sherbourne, Barbados.

Watson, P.K. and R. Clarke, 1997, A policy oriented macroeconometric model of Trinidad and Tobago, in: E. Edinval, A. Maurin and J.G. Montauban, eds., Stratégies de développement comparées dans la Caraïbe, L'Hermès, 265-337.

Watson, P.K. and S.S. Teelucksingh, 1997, Macroeconomic modelling and cointegration in a small sample framework (with application to Trinidad and Tobago), in: D. Worrell and R. Craigwell, eds., Macroeconomics and finance in the Caribbean: Quantitative analyses (Caribbean Centre for Monetary Studies, University of the West Indies) 13−51.

Watson, P.K. and S.S. Teelucksingh, 2000, Forecasting the Trinidad and Tobago macroeconomy using a medium-size structural econometric model: a policy making perspective", in: A. Maurin and P. Watson eds., Empirical Studies in Caribbean Economy (Caribbean Centre for Monetary Studies, University of the West Indies) 267−324.

Worrell, D., D. Marshall and N. Smith, 2000, The political economy of exchange rate policy in the Caribbean, Research Network Working paper R-401 (Inter-American Development Bank, Washington, DC).

8 Appendix: Tables and Figures

Table 1: An application of the test regression of Hylleberg *et al.* (1990)
Computed with a constant, trend and seasonal dummies

Null	m	y	$w1$	p	$p*$	pr	$m-p**$	$w1-p$
$t\pi_1$	-1.73	-2.78	-1.91	-0.26	-2.01	-0.35	-1.73	-2.04
$t\pi_2$	-4.22**	-4.23**	-3.99**	-4.10**	-4.69**	-4.60**	-4.21**	-4.79**
$F(\pi_3, \pi_4)$	26.77**	13.98**	27.45**	37.95**	29.79**	22.76**	26.19**	19.72**
$F(\pi_2, \pi_3, \pi_4)$	23.01**	15.55**	26.10**	28.49**	36.69**	40.36**	22.57**	19.77**
$F(\pi_1, ..., \pi_4)$	17.45**	11.76**	21.23**	21.79**	27.59**	30.96**	17.06**	16.36**
Sample	97:1-11:4	98:1-11:4	97:1-11:4	97:1-11:4	97:1-11:4	97:1-11:4	97:1-11:4	97:1-11:4
Lags	1	1	1	3	2	4	1	1
Dummies	-	11:3, 11:4	04:1	08:3, 10:2, 10:4	08:2, 08:4	08:3, 08:4, 10:2, 10:4	-	-

Table 2: An application of the test regression of Osborn *et al.* (1988)
Computed with a constant, seasonal dummies and trend

Null	m	y	$w1$	p	$p*$	pr	$m-p**$	$w1-p$
$t\pi_1$	-2.44	-4.41**	-0.36	-0.58	1.62	0.27	-2.57	-0.49
$t\pi_2$	-7.45**	-5.97**	-8.62**	-8.84**	-9.40**	-9.77**	-7.38**	-7.55**
$F(\pi_1, \pi_2)$	44.30**	40.35**	51.92**	79.43**	54.76**	72.16**	44.33**	41.35**
Sample	97:1-11:4	98:1-11:4	97:2-11:4	97:1-11:4	97:1-11:4	97:1-11:4	97:1-11:4	97:1-11:4
Lags	1	0	1	1	2	3	1	0
Dummies	-	11:3, 11:4	04:1	08:3, 10:2, 10:4	08:2, 08:4	08:4, 10:2, 10:4	-	-

Table 3: An application of the test regression of Osborn *et al.* (1988)
Computed with a constant, seasonal dummies, trend and seasonally trend dummies

$Null$	m	y	$w1$	p	$p*$	pr	$m-p**$	$w1-p$
$t\pi_1$	-2.17 *	-3.25**	-0.14	-0.01	2.22	0.33	-2.31*	-0.56
$t\pi_2$	-7.91**	-5.10**	-8.99**	-8.99**	-9.78**	-10.2**	-7.85**	-9.28**
$F(\pi_1, \pi_2)$	48.30**	24.64**	55.53**	77.55**	58.32**	76.68**	48.43**	61.46**
$Sample$	97:1-11:4	98:1-11:4	97:2-11:4	97:1-11:4	97:1-11:4	97:1-11:4	97:1-11:4	97:1-11:4
$Lags$	1	0	1	1	3	3	1	1
$Dummies$	-	-	04:1	08:3, 10:2, 10:4	08:2, 08:4	08:4, 10:2, 10:4	-	04:1

Table 4: F and related statistics for the sequential reduction from a fifth to first-order VAR
The sample is 1997(2)−2011(4) for 59 observations
Computed with a constant, trend, centered seasonal dummies and regime dummies Dc, Dg and Ds.

Null				Maintained			
$System$	k	λ	$SC/HQ/AIC$	$VAR(5)$	$VAR(4)$	$VAR(3)$	$VAR(2)$
VAR(5)	112	583.97	-12.06/-14.46/-15.99				
↓							
VAR(4)	96	568.43	-12.63/-14.70/-16.01	1.01 [0.45] (16,86)			
↓							
VAR(3)	80	550.49	-13.13/-14.85/-15.95	1.18 [0.26] (32,104)	1.35 [0.18] (16,98)		
↓							
VAR(2)	64	536.73	-13.77/-15.15/-16.03	1.18 [0.24] (48,109)	1.26 [0.18] (32,119)	1.14 [0.33] (16,110)	
↓							
VAR(1)	48	531.15	-14.69/-15.72/-16.38	1.01 [0.47] (64,111)	1.01 [0.46] (48,125)	0.82 [0.74] (32,134)	0.49 [0.95] (16,122)

Table 5: Cointegration analysis of time series data
The sample is 1997(2)-2011(4) for 59 observations
Computed with a constant, trend, centered seasonal dummies and regime dummies (D_c, D_g and D_s). All deterministic terms except the trend are set unrestricted in the cointegration space.

$Hypotheses$	$r = 1$	$r = 2$	$r = 3$	$r = 4$
$Eigenvalues$	0.462	0.410	0.356	0.082
λ_{trace}	98.82** [0.00]	62.22** [0.00]	31.05** [0.01]	5.06 [0.60]
λ^a_{trace}	92.12** [0.00]	58.00** [0.00]	28.95* [0.02]	4.72 [0.64]
λ_{max}	36.60* [0.01]	31.17** [0.01]	25.99** [0.00]	5.06 [0.60]
λ^a_{max}	34.12* [0.02]	29.05* [0.02]	24.23** [0.01]	4.72 [0.64]
$Eigenvectors\ \beta$				
$m - p*$	1	0.075	-0.302	-0.224
y	-0.913	1	-0.457	-0.442
$w1 - p$	-0.307	-0.027	1	-0.525
pr	1.228	-0.069	1.239	1
t	0.021	-0.023	0.010	0.038
$Weights\ \alpha$				
$m - p*$	-0.362	0.364	0.251	0.128
y	0.016	-0.697	0.046	0.015
$w1 - p$	0.081	0.034	-0.194	0.034
pr	-0.037	-0.053	-0.068	-0.005

Table 6: Structural hypotheses in the cointegration space

Hypotheses	Eigenvector(s)	Likelihood ratio (df)	p-value	$\hat{i} = a, b, c$
H_1: (1, -a, -b, 4b, ac)	1	0.008 (1)	0.929	$a = 0.874$, $b = 0.299$
H_2: (0, 1, 0, 0, -c)	2	0.011 (2)	0.994	$c = 0.023$
H_3: $H_1 + H_2$	1, 2	0.016 (3)	0.999	–
H_{a1}: (1, 0, -c, 4c, 0)	1	0.001 (2)	0.999	$c = 0.302$
H_{a2}: (1, -*, -c, 4c, **)	1	0.008 (1)	0.929	$c = 0.306$, $^* = 0.885$, $^{**} = 0.021$

Table 7: Long-run weak exogeneity tests (tested jointly with hypothesis H_3)

Hypotheses	Feedback Vectors	Likelihood ratio (df)	p-value	$\hat{\alpha}'$
H_{4a}: (*, 0, 0, 0)	1	19.154** (6)	0.004	(-0.49, 0, 0, 0)
H_{4b}: (*, 0, *, 0)	1	13.344* (5)	0.020	(-0.45, 0, 0.10, 0)
H_{4c}: (*, 0, 0, *)	1	6.653 (5)	0.248	(-0.36, 0, 0, -0.04)
H_{4d}: (*, *, 0, 0)	1	17.884** (5)	0.003	(-0.50, -0.05, 0, 0)
H_5: (0, *, 0, 0)	2	2.776 (6)	0.836	(0, -0.74, 0, 0)
H_6; $H_{4c} + H_5$	1, 2	9.543 (8)	0.299	–

Table 8: An unrestricted VEqCM of consumers' imports in TT 1997(2)-2011(4): Eq'n d(m-p*)

Variable	lag(j)				
	0	1	2	3	4
d(m-p*)	-	-0.292	-0.083	-0.195	-0.185
		(0.226)	(0.210)	(0.188)	(0.145)
dy	-	0.512	0.659	0.196	0.603
		(0.746)	(0.623)	(0.559)	(0.421)
d(w1-p)	-	0.150	0.279	-1.031	0.051
		(0.371)	(0.350)	(0.330)	(0.375)
dp*	-	0.071	5.939	0.962	-0.828
		(2.233)	(2.551)	(2.682)	(2.671)
dp	-	-0.108	1.367	0.395	1.344
		(1.583)	(1.264)	(1.312)	(1.136)
ecm1	-	-0.133	-	-	-
		(0.185)			
ecm2	-	0.430	-	-	-
		(0.826)			
Dc	-0.064	-	-	-	-
	(0.068)				
Dg	0.050	-	-	-	-
	(0.071)				
Ds	0.020	-	-	-	-
	(0.079)				
constant	-3.323	-	-	-	-
	(3.210)				
CS_t	-0.291	0.035	0.017	-	-
	(0.065)	(0.081)	(0.072)		

Table 9: An unrestricted VEqCM of consumers' imports in TT 1997(2)-2011(4): Eq'n dy

Variable	lag(j)				
	0	1	2	3	4
d(m-p)	-	0.161	0.070	0.015	0.022
		(0.086)	(0.079)	(0.072)	(0.055)
dy	-	0.028	-0.173	-0.390	-0.049
		(0.283)	(0.236)	(0.212)	(0.160)
d(w1-p)	-	0.013	0.073	0.074	0.332
		(0.141)	(0.133)	(0.125)	(0.142)
dp*	-	-1.072	0.807	-1.929	0.048
		(0.848)	(0.968)	(1.018)	(1.014)
dp	-	-0.109	-0.094	-0.734	0.394
		(0.601)	(0.480)	(0.498)	(0.431)
ecm1	-	-0.060	-	-	-
		(0.070)			
ecm2	-	-0.781	-	-	-
		(0.313)			
Dc	-0.021	-	-	-	-
	(0.026)				
Dg	0.037	-	-	-	-
	(0.027)				
Ds	-0.161	-	-	-	-
	(0.030)				
constant	2.122	-	-	-	-
	(1.218)				
CS_t	0.005	0.070	0.018	-	-
	(0.025)	(0.031)	(0.027)		

Table 10: An unrestricted VEqCM of consumers' imports in TT 1997(2)-2011(4): Eq'n d(w1-p)

Variable	lag(j)				
	0	1	2	3	4
d(m-p)	-	0.039	0.122	-0.012	0.011
		(0.086)	(0.080)	(0.072)	(0.055)
dy	-	-0.066	0.051	-0.157	-0.054
		(0.283)	(0.236)	(0.212)	(0.160)
d(w1-p)	-	0.161	0.039	0.076	-0.216
		(0.141)	(0.133)	(0.125)	(0.142)
dp*	-	-1.045	-0.921	1.576	-2.473
		(0.847)	(0.968)	(1.018)	(1.014)
dp	-	1.115	-0.549	0.733	0.143
		(601)	(0.480)	(0.498)	(0.431)
ecm1	-	0.010	-	-	-
		(0.070)			
ecm2	-	-0.020	-	-	-
		(0.313)			
Dc	-0.078	-	-	-	-
	(0.026)				
Dg	0.117	-	-	-	-
	(0.027)				
Ds	-0.032	-	-	-	-
	(0.030)				
constant	0.216	-	-	-	-
	(1.218)				
CS_t	0.054	0.017	0.056	-	-
	(0.025)	(0.031)	(0.027)		

Table 11: An unrestricted VEqCM of consumers' imports in TT 1997(2)-2011(4): Eq'n d(pr)

Variable	lag(j)				
	0	1	2	3	4
d(m-p)	-	0.046	0.011	0.004	-0.002
		(0.025)	(0.023)	(0.021)	(0.016)
dy	-	-0.129	-0.174	-0.068	-0.038
		(0.083)	(0.069)	(0.062)	(0.047)
d(w1-p)	-	-0.008	-0.002	-0.023	-0.034
		(0.041)	(0.039)	(0.037)	(0.042)
dp*	-	-0.328	-0.264	-0.382	-0.195
		(0.248)	(0.283)	(0.298)	(0.296)
dp	-	0.199	-0.011	-0.021	0.053
		(0.176)	(0.140)	(0.146)	(0.126)
ecm1	-	-0.060	-	-	-
		(0.021)			
ecm2	-	0.050	-	-	-
		(0.092)			
Dc	-0.054	-	-	-	-
	(0.008)				
Dg	-0.005	-	-	-	-
	(0.008)				
Ds	-0.012	-	-	-	-
	(0.009)				
constant	-0.919	-	-	-	-
	(0.356)				
CS_t	0.005	0.009	-0.010	-	-
	(0.007)	(0.009)	(0.008)		

Table 12: Correlation of structural residuals (standard deviations on diagonal)

	d(m-p*)	dy	d(w1-p)	dpr
d(m-p*)	0.081	0.127	0.193	-0.060
dy	0.127	0.031	0.105	-0.122
d(w1-p)	0.193	0.105	0.031	0.033
dpr	-0.060	-0.122	0.033	0.009

Table 13: Correlation of structural residuals (standard deviations on diagonal)

	d(m-p*)	dy	d(w1-p)	dpr
d(m-p*)	0.071	0.093	0.234	-0.092
dy	0.093	0.030	0.054	-0.195
d(w1-p)	0.234	0.054	0.029	0.083
dpr	-0.092	-0.195	0.083	0.008

Table 14: Single-equation and system diagnostics using reduced-form residuals

Eq'n	Test	Test type	Stat [p-value]
d(m-p*)	AR 1-4	F(4,43)	0.499 [0.736]
d(m-p*)	ARCH 1-4	F(4,51)	0.639 [0.638]
d(m-p*)	Normality	$\chi^2(2)$	1.418 [0.492]
d(m-p*)	Hetero	F(19,39)	1.013 [0.469]
dy	AR 1-4	F(4,48)	1.206 [0.321]
dy	ARCH 1-4	F(4,51)	0.545 [0.703]
dy	Normality	$\chi^2(2)$	2.080 [0.354]
dy	Hetero	F(8,50)	1.078 [0.394]
d(w1-p)	AR 1-4	F(4,46)	0.763 [0.555]
d(w1-p)	ARCH 1-4	F(4,51)	0.808 [0.526]
d(w1-p)	Normality	$\chi^2(2)$	0.689 [0.708]
d(w1-p)	Hetero	F(12,46)	1.990* [0.048]
dpr	AR 1-4	F(4,45)	0.455 [0.768]
dpr	ARCH 1-4	F(4,51)	0.358 [0.837]
dpr	Normality	$\chi^2(2)$	5.913 [0.052]
dpr	Hetero	F(14,44)	0.454 [0.945]
Vector	SEM-AR 1-4	F(64,123)	0.839 [0.780]
Vector	Normality	$\chi^2(8)$	11.241 [0.188]

Figure 1: TT time series data, 1996(1)−2011(4)

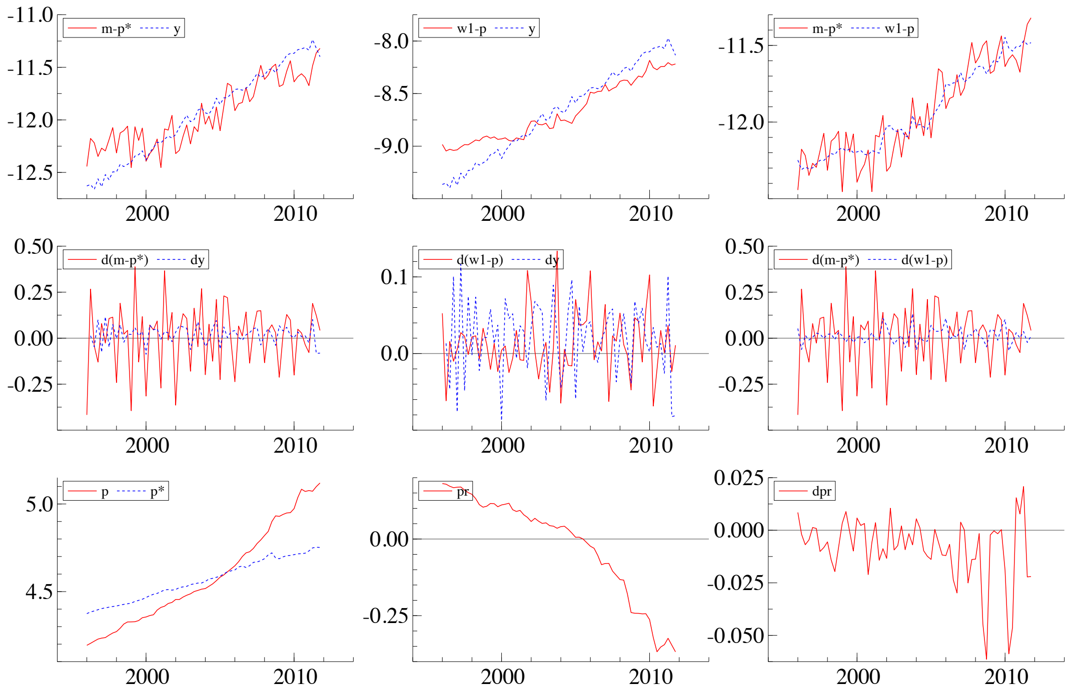

Table 15: Encompassing test statistics

Test	Distribution	WT v B	p-value	Distribution	B v WT	p-value
Cox	N(0,1)	-10.44**	0.000	N(0,1)	-2.127*	0.033
Ericsson IV	N(0,1)	7.338**	0.000	N(0,1)	1.843	0.065
Sargan	$\chi^2(7)$	23.26**	0.001	$\chi^2(2)$	2.637	0.268
Joint Model	F(7,46)	6.127**	0.001	F(2,46)	1.337	0.273
Test	Distribution	M v B	p-value	Distribution	B v M	p-value
Cox	N(0,1)	-9.616**	0.000	N(0,1)	-1.691	0.091
Ericsson IV	N(0,1)	6.833**	0.000	N(0,1)	1.477	0.140
Sargan	$\chi^2(8)$	22.08**	0.001	$\chi^2(3)$	2.721	0.437
Joint Model	F(8,44)	5.658**	0.000	F(3,45)	0.902	0.448

Figure 2: ACFs and PACFs of log-levels and log first-differences, 1996(1)−2011(4)

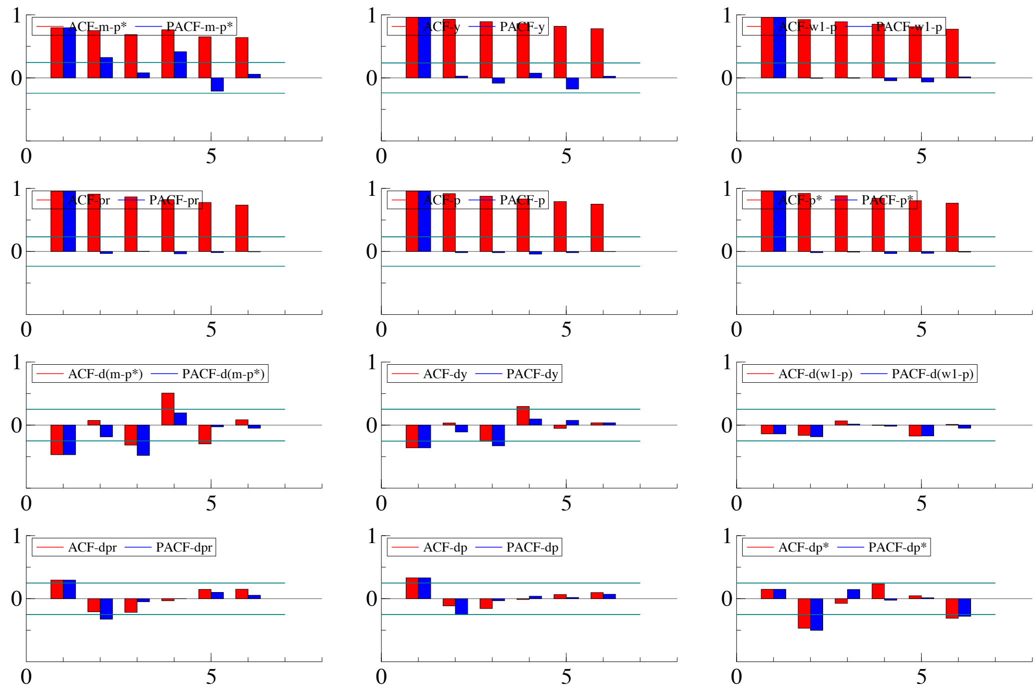

Figure 3: System graphical diagnostic information

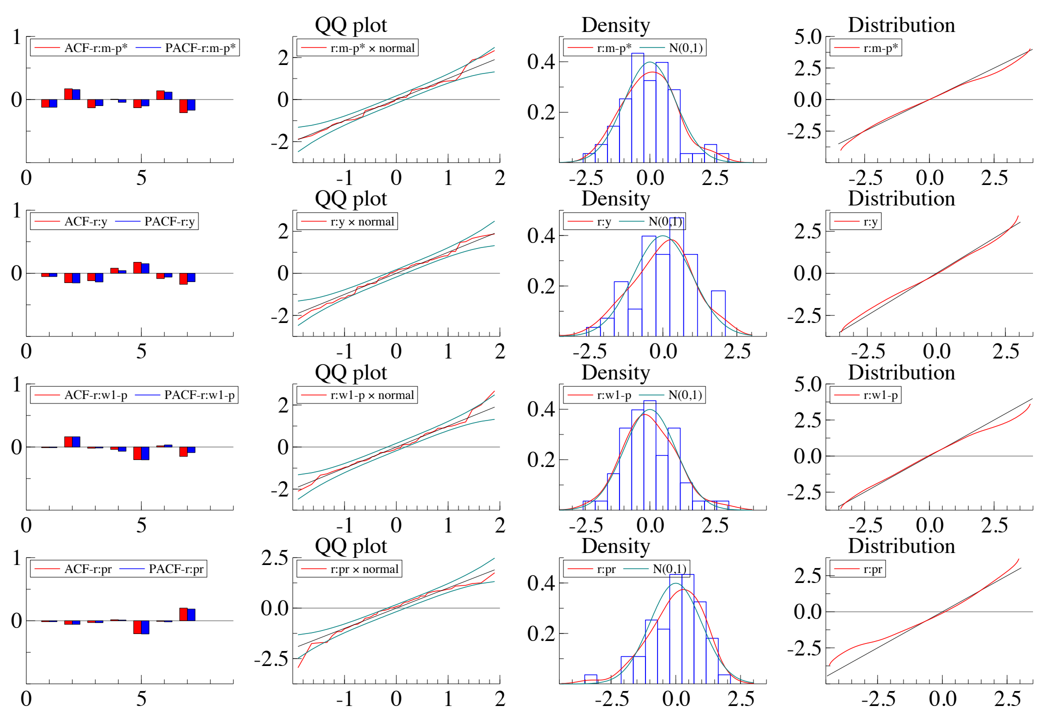

Figure 4: System recursive evaluation statistics

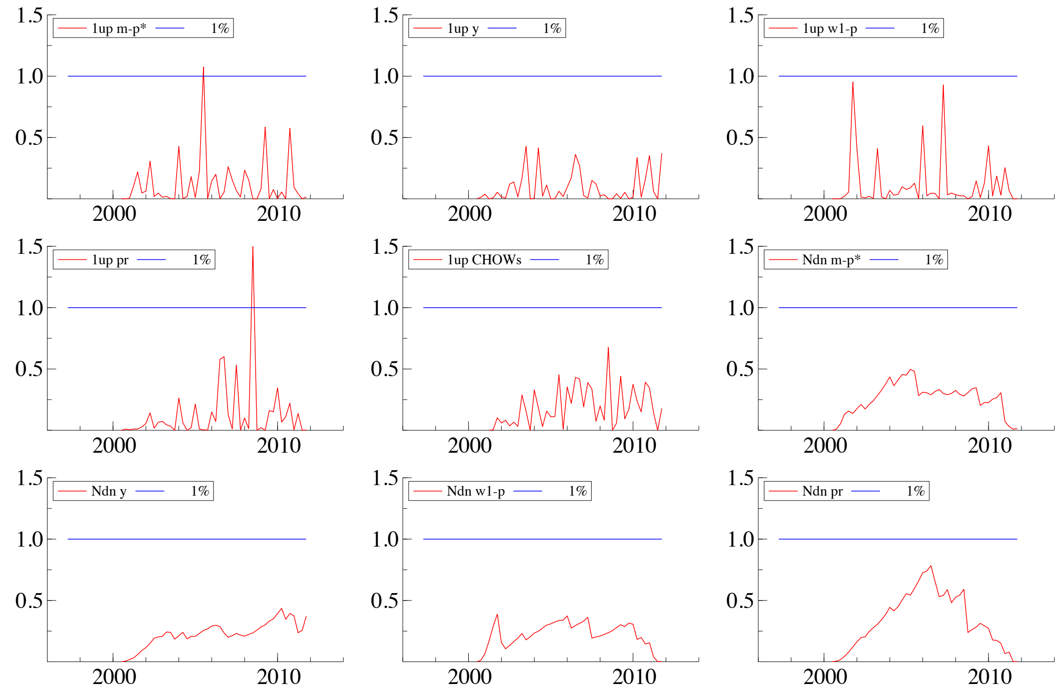

Figure 5: Recursive eigenvalues

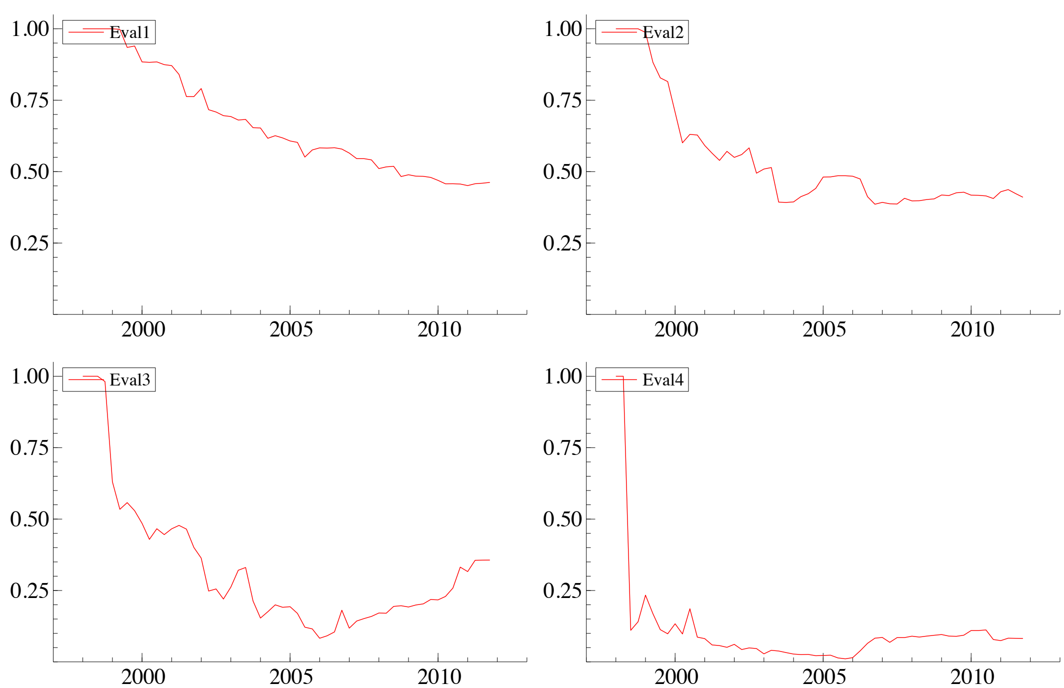

Figure 6: Long-run relation vectors

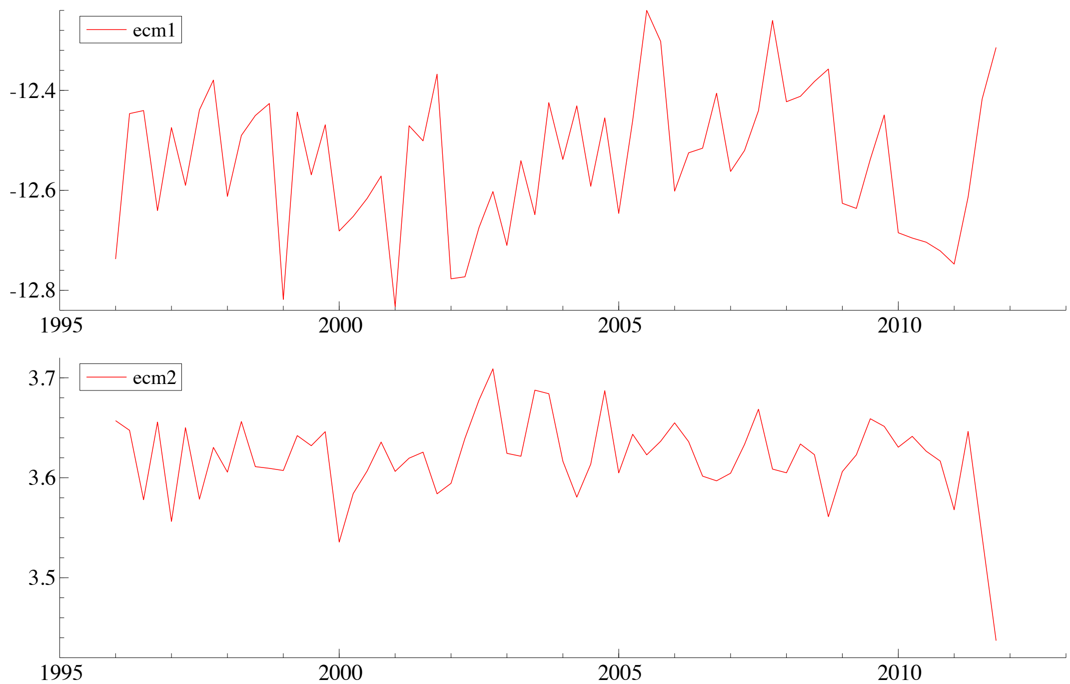

Figure 7: Graphical regression information for model with separated price effects

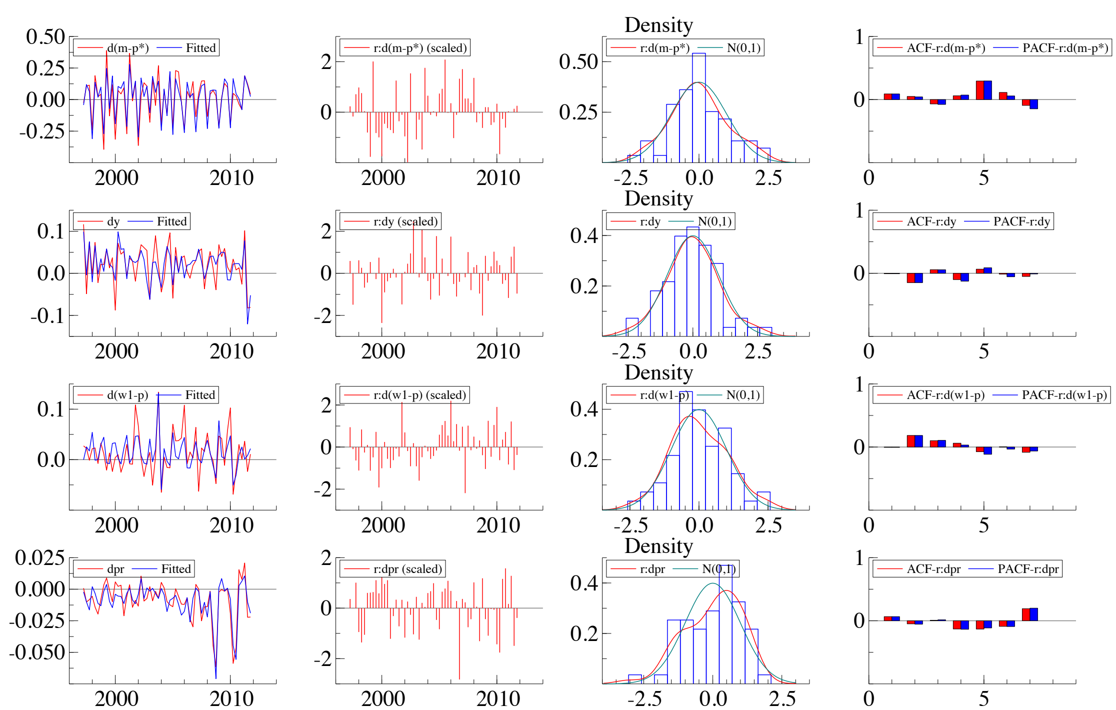

Figure 8: Recursive FIML statistics

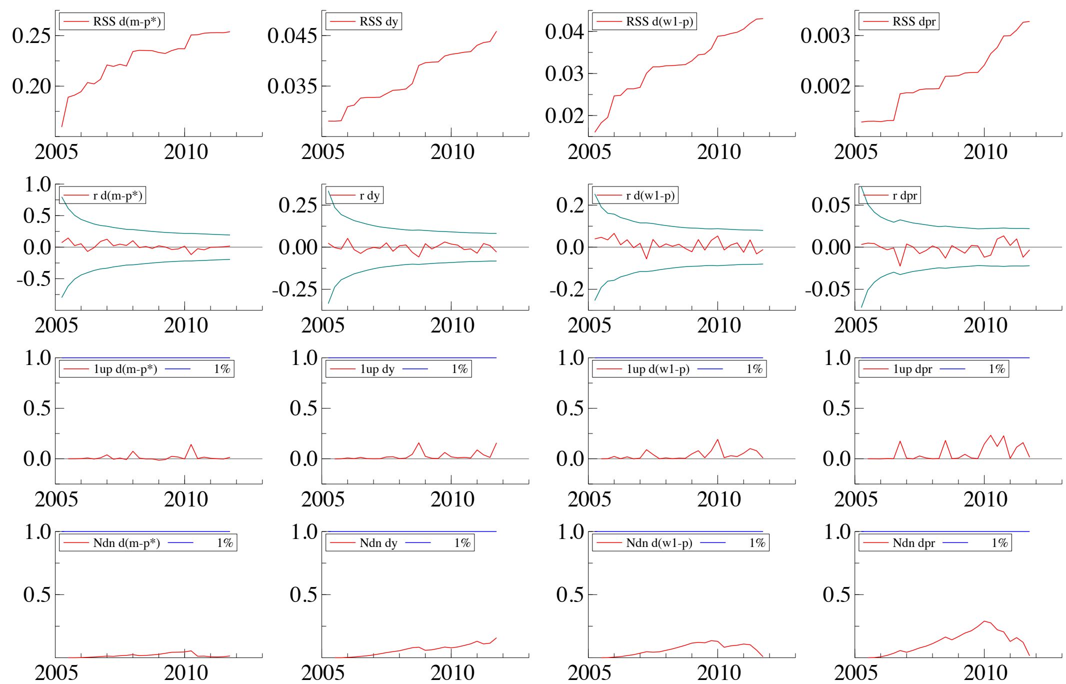

Figure 9: System 1-step ahead forecasts