

Short run and long run causality in time series: inference

Jean-Marie Dufour,^{*†} Denis Pelletier[‡] and Éric Renault[§]

Université de Montréal

September 2003

March 2005

ABSTRACT

We propose methods for testing hypothesis of non-causality at various horizons, as defined in Dufour and Renault (1998, *Econometrica*). We study in detail the case of VAR models and we propose linear methods based on running vector autoregressions at different horizons. While the hypotheses considered are nonlinear, the proposed methods only require linear regression techniques as well as standard Gaussian asymptotic distributional theory. Bootstrap procedures are also considered. For the case of integrated processes, we propose extended regression methods that avoid nonstandard asymptotics. The methods are applied to a VAR model of the U.S. economy.

Key words: Granger causality; indirect causality; vector autoregression; bootstrap; macroeconomics.

JEL classification numbers: C1; C12; C15; C32; C51; C53; E3; E4; E52.

Forthcoming in the *Journal of Econometrics*.

^{*} CIRANO, CIREQ, and Département de sciences économiques, Université de Montréal, C.P. 6128 succursale Centre-ville, Montréal, Québec, Canada H3C 3J7

[†] Corresponding author. TEL: 1-514-343-2400; FAX: 1-514-343-5831.
E-mail: jean.marie.dufour@umontreal.ca (Jean-Marie Dufour).
Web page: <http://www.fas.umontreal.ca/SCECO/Dufour>

[‡] CIRANO, CIREQ, Université de Montréal, and Department of Economics, North Carolina State University, Campus Box 8110, Raleigh, NC, USA 27695-8110

[§] CIRANO, CIREQ, and Département de sciences économiques, Université de Montréal, C.P. 6128 succursale Centre-ville, Montréal, Québec, Canada H3C 3J7

1. Introduction

The concept of causality introduced by Wiener (1956) and Granger (1969) is now a basic notion for studying dynamic relationships between time series. The literature on this topic is considerable; see, for example, the reviews of Pierce and Haugh (1977), Newbold (1982), Geweke (1984), Lütkepohl (1991) and Gouriéroux and Monfort (1997, Chapter 10). The original definition of Granger (1969), which is used or adapted by most authors on this topic, refers to the predictability of a variable $X(t)$, where t is an integer, from its own past, the one of another variable $Y(t)$ and possibly a vector $Z(t)$ of auxiliary variables, **one period ahead**: more precisely, we say that Y causes X in the sense of Granger if the observation of Y up to time t ($Y(\tau) : \tau \leq t$) can help one to predict $X(t + 1)$ when the corresponding observations on X and Z are available ($X(\tau), Z(\tau) : \tau \leq t$); a more formal definition will be given below.

Recently, however, Lütkepohl (1993) and Dufour and Renault (1998) have noted that, for multivariate models where a vector of auxiliary variables Z is used in addition to the variables of interest X and Y , it is possible that Y does not cause X in this sense, but can still help to predict X **several periods ahead**; on this issue, see also Sims (1980), Renault, Sekkat and Szafarz (1998) and Giles (2002). For example, the values $Y(\tau)$ up to time t may help to predict $X(t + 2)$, even though they are useless to predict $X(t + 1)$. This is due to the fact that Y may help to predict Z one period ahead, which in turn has an effect on X at a subsequent period. It is clear that studying such indirect effects can have a great interest for analyzing the relationships between time series. In particular, one can distinguish in this way properties of “short-run (non-)causality” and “long-run (non-)causality”.

In this paper, we study the problem of testing non-causality at various horizons as defined in Dufour and Renault (1998) for finite-order vector autoregressive (VAR) models. In such models, the non-causality restriction at horizon one takes the form of relatively simple zero restrictions on the coefficients of the VAR [see Boudjellaba, Dufour and Roy (1992) and Dufour and Renault (1998)]. However non-causality restrictions at higher horizons (greater than or equal to 2) are generally nonlinear, taking the form of zero restrictions on multilinear forms in the coefficients of the VAR. When applying standard test statistics such as Wald-type test criteria, such forms can easily lead to asymptotically singular covariance matrices, so that standard asymptotic theory would not apply to such statistics. Further, calculation of the relevant covariance matrices _ which involve the derivatives of potentially large numbers of restrictions _ can become quite awkward.

Consequently, we propose simple tests for non-causality restrictions at various horizons [as defined in Dufour and Renault (1998)] which can be implemented only through linear regression methods and do not involve the use of artificial simulations [e.g., as in Lütkepohl and Burda (1997)]. This will be done, in particular, by considering multiple horizon vector autoregressions [called (p, h) -autoregressions] where the parameters of interest can be estimated by linear methods. Restrictions of non-causality at different horizons may then be tested through simple Wald-type (or Fisher-type) criteria after taking into account the fact that such autoregressions involve autocorrelated errors [following simple moving average processes] which are orthogonal to the regressors. The correction for the presence of autocorrelation in the errors may then be performed by using an autocorrelation consistent [or heteroskedasticity-autocorrelation-consistent (HAC)] covariance matrix estimator. Further, we distinguish between the case where the VAR process considered is stable (*i.e.*, the roots of the determinant of the associated AR polynomial are all outside the unit circle) and the one where the process may be integrated of an unknown order (although not explosive). In the first case, the test statistics follow standard chi-square distributions while, in the second case, they may follow nonstandard asymptotic distributions involving nuisance parameters, as already observed by several authors for the case of causality tests at horizon one [see Sims, Stock and Watson (1990), Toda and Phillips (1993, 1994), Toda and Yamamoto (1995), Dolado and Lütkepohl (1996) and Yamada and Toda (1998)]. To meet the objective of producing simple procedures that can be implemented by least squares methods, we propose to deal with such problems by using an extension to the case of multiple horizon autoregressions of the lag extension technique suggested by Choi (1993) for inference on univariate autoregressive models and by Toda and Yamamoto (1995) and Dolado and Lütkepohl (1996) for inference on standard VAR models. This extension will allow us to use standard asymptotic theory in order to test non-causality at different horizons without making assumption on the presence of unit roots and cointegrating relations. Finally, to alleviate the problems of finite-sample unreliability of asymptotic approximations in VAR models (on both stationary and nonstationary series), we propose the use of bootstrap methods to implement the proposed test statistics.

In section 2, we describe the model considered and introduce the notion of autoregression at horizon h [or (p, h) -autoregression] which will be the basis of our method. In section 3, we study the estimation of (p, h) -autoregressions and the asymptotic distribution of the relevant estimators for stable VAR processes. In section 4, we study the testing of non-causality at various horizons for stationary processes, while in section 5, we consider the case of processes that may

be integrated. In section 6, we illustrate the procedures on a monthly VAR model of the U.S. economy involving a monetary variable (nonborrowed reserves), an interest rate (federal funds rate), prices (GDP deflator) and real GDP, over the period 1965-1996. We conclude in section 7.

2. Multiple horizon autoregressions

In this section, we develop the notion of “autoregression at horizon h ” and the relevant notations. Consider a VAR (p) process of the form:

$$W(t) = \mu(t) + \sum_{k=1}^p \pi_k W(t-k) + a(t), \quad t = 1, \dots, T, \quad (1)$$

where $W(t) = (w_{1t}, w_{2t}, \dots, w_{mt})'$ is an $m \times 1$ random vector, $\mu(t)$ is a deterministic trend, and

$$\begin{aligned} \mathbb{E}[a(s) a(t)'] &= \Omega, \text{ if } s = t, \\ &= 0, \text{ if } s \neq t, \end{aligned} \quad (2)$$

$$\det(\Omega) \neq 0. \quad (3)$$

The most common specification for $\mu(t)$ consists in assuming that $\mu(t)$ is a constant vector, *i.e.*

$$\mu(t) = \mu, \quad (4)$$

although other deterministic trends could also be considered.

The VAR (p) in equation (1) is an autoregression at horizon 1. We can then also write for the observation at time $t + h$:

$$W(t+h) = \mu^{(h)}(t) + \sum_{k=1}^p \pi_k^{(h)} W(t+1-k) + \sum_{j=0}^{h-1} \psi_j a(t+h-j), \quad t = 0, \dots, T-h,$$

where $\psi_0 = I_m$ and $h < T$. The appropriate formulas for the coefficients $\pi_k^{(h)}$, $\mu^{(h)}(t)$ and ψ_j are given in Dufour and Renault (1998), namely:

$$\pi_k^{(h+1)} = \pi_{k+h} + \sum_{l=1}^h \pi_{h-l+1} \pi_k^{(l)} = \pi_{k+1}^{(h)} + \pi_1^{(h)} \pi_k, \quad \pi_1^{(0)} = I_m, \quad \pi_k^{(1)} = \pi_k, \quad (5)$$

$$\mu^{(h)}(t) = \sum_{k=0}^{h-1} \pi_1^{(k)} \mu(t+h-k), \quad \psi_h = \pi_1^{(h)}, \quad \forall h \geq 0. \quad (6)$$

The ψ_h matrices are the impulse response coefficients of the process, which can also be obtained from the formal series:

$$\psi(z) = \pi(z)^{-1} = I_m + \sum_{k=1}^{\infty} \psi_k z^k, \quad \pi(z) = I_m - \sum_{k=1}^{\infty} \pi_k z^k. \quad (7)$$

Equivalently, the above equation for $W(t+h)$ can be written in the following way:

$$\begin{aligned} W(t+h)' &= \mu^{(h)}(t)' + \sum_{k=1}^p W(t+1-k)' \pi_k^{(h)'} + u^{(h)}(t+h)' \\ &= \mu^{(h)}(t)' + W(t,p)' \pi^{(h)} + u^{(h)}(t+h)', \quad t = 0, \dots, T-h, \end{aligned} \quad (8)$$

where $W(t,p)' = [W(t)', W(t-1)', \dots, W(t-p+1)']$, $\pi^{(h)} = [\pi_1^{(h)}, \dots, \pi_p^{(h)}]'$ and

$$u^{(h)}(t+h)' = [u_1^{(h)}(t+h), \dots, u_m^{(h)}(t+h)] = \sum_{j=0}^{h-1} a(t+h-j)' \psi_j'.$$

It is straightforward to see that $u^{(h)}(t+h)$ has a non-singular covariance matrix.

We call (8) an “autoregression of order p at horizon h ” or a “ (p, h) -autoregression”. In the sequel, we will assume that the deterministic part of each autoregression is a linear function of a finite-dimensional parameter vector, *i.e.*

$$\mu^{(h)}(t) = \gamma(h) D^{(h)}(t) \quad (9)$$

where $\gamma(h)$ is a $m \times n$ coefficient vector and $D^{(h)}(t)$ is a $n \times 1$ vector of deterministic regressors. If $\mu(t)$ is a constant vector, *i.e.* $\mu(t) = \mu$, then $\mu^{(h)}(t)$ is simply a constant vector (which may depend on h):

$$\mu^{(h)}(t) = \mu_h. \quad (10)$$

To derive inference procedures, it will be useful to put (8) in matrix form, which yields:

$$w_h(h) = \overline{W}_p(h) \Pi^{(h)} + U_h(h), \quad h = 1, \dots, H, \quad (11)$$

where $w_h(k)$ and $U_h(k)$ are $(T-k+1) \times m$ matrices and $\overline{W}_p(k)$ is a $(T-k+1) \times (n+mp)$

matrix defined as

$$w_h(k) = \begin{bmatrix} W(0+h)' \\ W(1+h)' \\ \vdots \\ W(T-k+h)' \end{bmatrix} = [w_1(h, k), \dots, w_m(h, k)], \quad (12)$$

$$\overline{W}_p(k) = \begin{bmatrix} W_p(0)' \\ W_p(1)' \\ \vdots \\ W_p(T-k)' \end{bmatrix}, \quad W_p(t) = \begin{bmatrix} D^{(h)}(t)' \\ W(t, p) \end{bmatrix}, \quad (13)$$

$$\Pi^{(h)} = \begin{bmatrix} \gamma^{(h)'} \\ \pi^{(h)} \end{bmatrix} = [\beta_1(h), \beta_2(h), \dots, \beta_m(h)], \quad (14)$$

$$U_h(k) = \begin{bmatrix} u^{(h)}(0+h)' \\ u^{(h)}(1+h)' \\ \vdots \\ u^{(h)}(T-k+h)' \end{bmatrix} = [u_1(h, k), \dots, u_m(h, k)], \quad (15)$$

$$u_i(h, k) = [u_i^{(h)}(0+h), u_i^{(h)}(1+h), \dots, u_i^{(h)}(T-k+h)]'. \quad (16)$$

We shall call the formulation (11) a “ (p, h) -autoregression in matrix form”. Other formulations could also be written by stacking autoregressions at different horizons; see the discussion paper version of this article [Dufour, Pelletier and Renault (2003)].

3. Estimation of (p, h) autoregressions

Let us now consider each autoregression of order p at horizon h as given by (11):

$$w_h(h) = \overline{W}_p(h) \Pi^{(h)} + U_h(h), \quad h = 1, \dots, H. \quad (17)$$

We can estimate (17) by ordinary least squares (OLS), which yields the estimator:

$$\hat{\Pi}^{(h)} = [\overline{W}_p(h)' \overline{W}_p(h)]^{-1} \overline{W}_p(h)' w_h(h) = \Pi^{(h)} + [\overline{W}_p(h)' \overline{W}_p(h)]^{-1} \overline{W}_p(h)' U_h(h),$$

hence

$$\sqrt{T} [\hat{\Pi}^{(h)} - \Pi^{(h)}] = \left[\frac{1}{T} \overline{W}_p(h)' \overline{W}_p(h) \right]^{-1} \frac{1}{\sqrt{T}} \overline{W}_p(h)' U_h(h)$$

where

$$\frac{1}{T} \overline{W}_p(h)' \overline{W}_p(h) = \frac{1}{T} \sum_{t=0}^{T-h} W_p(t) W_p(t)', \quad \frac{1}{\sqrt{T}} \overline{W}_p(h)' U_h(h) = \frac{1}{\sqrt{T}} \sum_{t=0}^{T-h} W_p(t) u^{(h)}(t+h)' .$$

Suppose now that

$$\frac{1}{T} \sum_{t=0}^{T-h} W_p(t) W_p(t)' \xrightarrow[T \rightarrow \infty]{P} \Gamma_p \quad \text{with} \quad \det(\Gamma_p) \neq 0. \quad (18)$$

In particular, this will be the case if the process $W(t)$ is second-order stationary, strictly indeter-
ministic and regular, in which case

$$\mathbb{E} [W_p(t) W_p(t)'] = \Gamma_p, \quad \forall t. \quad (19)$$

Cases where the process does not satisfy these conditions are covered in section 5. Further, since

$$u^{(h)}(t+h) = a(t+h) + \sum_{k=1}^{h-1} \psi_k a(t+h-k)$$

(where, by convention, any sum of the form $\sum_{k=1}^{h-1}$ with $h < 2$ is zero), we have:

$$\begin{aligned} \mathbb{E} [W_p(t) u^{(h)}(t+h)'] &= 0, \quad \text{for } h = 1, 2, \dots, \\ \vee \left\{ \text{vec} [W_p(t) u^{(h)}(t+h)'] \right\} &= \Delta_p(h). \end{aligned}$$

If the process $W(t)$ is strictly stationary with i.i.d. innovations $a(t)$ and finite fourth moments, we can write:

$$\mathbb{E}[W_p(s) u_i^{(h)}(s+h) u_j^{(h)}(t+h) W_p(t)'] = \Gamma_{ij}(p, h, t-s) = \Gamma_{ij}(p, h, s-t) \quad (20)$$

where $1 \leq i \leq m, 1 \leq j \leq m$, with

$$\begin{aligned} \Gamma_{ij}(p, h, 0) &= \mathbb{E} [W_p(t) u_i^{(h)}(t+h) u_j^{(h)}(t+h) W_p(t)'] \\ &= \sigma_{ij}(h) \mathbb{E} [W_p(t) W_p(t)'] = \sigma_{ij}(h) \Gamma_p, \end{aligned} \quad (21)$$

$$\Gamma_{ij}(p, h, t-s) = 0, \quad \text{if } |t-s| \geq h. \quad (22)$$

In this case,¹

$$\Delta_p(h) = [\sigma_{ij}(h) \Gamma_p]_{i,j=1,\dots,m} = \Sigma(h) \otimes \Gamma_p \quad (23)$$

where $\Sigma(h)$ is nonsingular, and thus $\Delta_p(h)$ is also nonsingular. The nonsingularity of $\Sigma(h)$ follows from the identity

$$u^{(h)}(t+h) = [\psi_{h-1}, \psi_{h-2}, \dots, \psi_1, I_m] [a(t+1)', a(t+2)', \dots, a(t+h)']'.$$

Under usual regularity conditions,

$$\frac{1}{\sqrt{T}} \sum_{t=0}^{T-h} \text{vec}[W_p(t) u^{(h)}(t+h)'] \xrightarrow[T \rightarrow \infty]{L} N[0, \bar{\Delta}_p(h)] \quad (24)$$

where $\bar{\Delta}_p(h)$ is a nonsingular covariance matrix which involves the variance and the autocovariances of $W_p(t) u^{(h)}(t+h)'$ [and possibly other parameters, if the process $W(t)$ is not linear].

Then,

$$\begin{aligned} \sqrt{T} \text{vec}[\hat{\Pi}^{(h)} - \Pi^{(h)}] &= \left\{ I_m \otimes \left[\frac{1}{T} \bar{W}_p(h)' \bar{W}_p(h) \right]^{-1} \right\} \text{vec} \left[\frac{1}{\sqrt{T}} \bar{W}_p(h)' U_h(h) \right] \\ &= \left\{ I_m \otimes \left[\frac{1}{T} \bar{W}_p(h)' \bar{W}_p(h) \right]^{-1} \right\} \frac{1}{\sqrt{T}} \sum_{t=0}^{T-h} \text{vec} [W_p(t) u^{(h)}(t+h)'] \\ &\xrightarrow[T \rightarrow \infty]{L} N[0, (I_m \otimes \Gamma_p^{-1}) \bar{\Delta}_p(h) (I_m \otimes \Gamma_p^{-1})]. \end{aligned} \quad (25)$$

For convenience, we shall summarize the above observations in the following proposition.

Proposition 1 ASYMPTOTIC NORMALITY OF LS IN A (p, h) STATIONARY VAR. *Under the assumptions (1), (18), and (24), the asymptotic distribution of $\sqrt{T} \text{vec}[\hat{\Pi}^{(h)} - \Pi^{(h)}]$ is $N[0, \Sigma(\hat{\Pi}^{(h)})]$, where $\Sigma(\hat{\Pi}^{(h)}) = (I_m \otimes \Gamma_p^{-1}) \bar{\Delta}_p(h) (I_m \otimes \Gamma_p^{-1})$.*

¹Note that (21) holds under the assumption of martingale difference sequence on $a(t)$. But to get (22) and allow the use of simpler central limit theorems, we maintain the stronger assumption that the innovations $a(t)$ are i.i.d. according to some distribution with finite fourth moments (not necessarily Gaussian).

4. Causality tests based on stationary (p, h) -autoregressions

Consider the i -th equation ($1 \leq i \leq m$) in system (11):

$$\bar{w}_i(h) = \overline{W}_p(h) \beta_i(h) + \bar{u}_i(h), \quad 1 \leq i \leq m, \quad (26)$$

where $\bar{w}_i(h) = w_i(h, h)$ and $\bar{u}_i(h) = u_i(h, h)$, where $w_i(h, h)$ and $u_i(h, h)$ are defined in (12) and (15). We wish to test:

$$H_0(h) : R\beta_i(h) = r \quad (27)$$

where R is a $q \times (n + mp)$ matrix of rank q . In particular, if we wish to test the hypothesis that w_{jt} does not cause w_{it} at horizon h [*i.e.*, using the notation of Dufour and Renault (1998), $w_j \not\rightarrow_h w_i | I_{(j)}$, where $I_{(j)}(t)$ is the Hilbert space generated by the basic information set $I(t)$ and the variables $w_{k\tau}$, $\omega < \tau \leq t$, $k \neq j$, ω being an appropriate starting time ($\omega \leq -p + 1$)], the restriction would take the form:

$$H_{j \rightarrow i}^{(h)} : \pi_{ijk}^{(h)} = 0, \quad k = 1, \dots, p, \quad (28)$$

where $\pi_k^{(h)} = \left[\pi_{ijk}^{(h)} \right]_{i,j=1,\dots,m}$, $k = 1, \dots, p$. In other words, the null hypothesis takes the form of a set of zero restrictions on the coefficients of $\beta_i(h)$ as defined in (14). The matrix of restrictions R in this case takes the form $R = R(j)$, where $R(j) \equiv [\delta_1(j), \delta_2(j), \dots, \delta_p(j)]'$ is a $p \times (n + mp)$ matrix, $\delta_k(j)$ is a $(n + pm) \times 1$ vector whose elements are all equal to zero except for a unit value at position $n + (k - 1)m + j$, *i.e.* $\delta_k(j) = [\delta(1, n + (k - 1)m + j), \dots, \delta(n + pm, n + (k - 1)m + j)]'$, $k = 1, \dots, p$, with $\delta(i, j) = 1$ if $i = j$, and $\delta(i, j) = 0$ if $i \neq j$. Note also that the conjunction of the hypothesis $H_{j \rightarrow i}^{(h)}$, $h = 1, \dots, (m - 2)p + 1$, is sufficient to obtain noncausality at all horizons [see Dufour and Renault (1998, section 4)]. Non-causality up to horizon H is the conjunction of the hypothesis $H_{j \rightarrow i}^{(h)}$, $h = 1, \dots, H$.

We have:

$$\hat{\beta}_i(h) = \beta_i(h) + [\overline{W}_p(h)' \overline{W}_p(h)]^{-1} \overline{W}_p(h)' \bar{u}_i(h),$$

hence

$$\sqrt{T} \left[\hat{\beta}_i(h) - \beta_i(h) \right] = \left[\frac{1}{T} \overline{W}_p(h)' \overline{W}_p(h) \right]^{-1} \frac{1}{\sqrt{T}} \sum_{t=0}^{T-h} \overline{W}_p(t) u_i^{(h)}(t+h).$$

Under standard regularity conditions [see White (1999, chap. 5-6)],

$$\sqrt{T} [\hat{\beta}_i(h) - \beta_i(h)] \xrightarrow[T \rightarrow \infty]{L} N[0, \mathbf{V}(\hat{\beta}_i)]$$

with $\det[\mathbf{V}(\hat{\beta}_i)] \neq 0$, where $\mathbf{V}(\hat{\beta}_i)$ can be consistently estimated:

$$\hat{V}_T(\hat{\beta}_i) \xrightarrow[T \rightarrow \infty]{P} \mathbf{V}(\hat{\beta}_i).$$

More explicit forms for $\hat{V}_T(\hat{\beta}_i)$ will be discussed below. Note also that

$$\Gamma_p = \text{plim}_{T \rightarrow \infty} \frac{1}{T} \overline{W}_p(h)' \overline{W}_p(h), \quad \det(\Gamma_p) \neq 0.$$

Let

$$\begin{aligned} V_{ip}(T) &= \text{Var} \left[\frac{1}{\sqrt{T}} \overline{W}_p(h)' \bar{u}_i(h) \right] = \frac{1}{T} \text{Var} \left[\sum_{t=0}^{T-h} W_p(t) u_i^{(h)}(t+h) \right] \\ &= \frac{1}{T} \left\{ \sum_{t=0}^{T-h} \mathbf{E} [W_p(t) u_i^{(h)}(t+h) u_i^{(h)}(t+h) W_p(t)'] \right. \\ &\quad + \sum_{\tau=1}^{h-1} \sum_{t=\tau+1}^{T-h} \left[\mathbf{E} [W_p(t) u_i^{(h)}(t+h) u_i^{(h)}(t-\tau+h) W_p(t-\tau)'] \right. \\ &\quad \left. \left. + \mathbf{E} [W_p(t-\tau) u_i^{(h)}(t-\tau+h) u_i^{(h)}(t+h) W_p(t)'] \right] \right\}. \end{aligned}$$

Let us assume that

$$V_{ip}(T) \xrightarrow[T \rightarrow \infty]{} V_{ip}, \quad \det(V_{ip}) \neq 0, \quad (29)$$

where V_{ip} can be estimated by a computable consistent estimator $\hat{V}_{ip}(T)$:

$$\hat{V}_{ip}(T) \xrightarrow[T \rightarrow \infty]{P} V_{ip}. \quad (30)$$

Then,

$$\sqrt{T} [\hat{\beta}_i(h) - \beta_i(h)] \xrightarrow[T \rightarrow \infty]{L} N[0, \Gamma_p^{-1} V_{ip} \Gamma_p^{-1}]$$

so that $\mathbf{V}(\hat{\beta}_i) = \Gamma_p^{-1} V_{ip} \Gamma_p^{-1}$. Further, in this case,

$$\hat{V}_T(\hat{\beta}_i) = \hat{\Gamma}_p^{-1} \hat{V}_{ip}(T) \hat{\Gamma}_p^{-1} \xrightarrow[T \rightarrow \infty]{P} \mathbf{V}(\hat{\beta}_i),$$

$$\hat{\Gamma}_p = \frac{1}{T} \sum_{t=0}^{T-h} W_p(t) W_p(t)' = \frac{1}{T} \overline{W}_p(h)' \overline{W}_p(h) \xrightarrow[T \rightarrow \infty]{\mathbf{p}} \Gamma_p.$$

We can thus state the following proposition.

Proposition 2 ASYMPTOTIC DISTRIBUTION OF TEST CRITERION FOR NON-CAUSALITY AT HORIZON h IN A STATIONARY VAR. *Suppose the assumptions of Proposition 1 hold jointly with (29) – (30). Then, under any hypothesis of the form $H_0(h)$ in (27), the asymptotic distribution of*

$$\mathcal{W}[H_0(h)] = T [R\hat{\beta}_i(h) - r]' [R\hat{V}_T(\hat{\beta}_i)R']^{-1} [R\hat{\beta}_i(h) - r] \quad (31)$$

is $\chi^2(q)$. In particular, under the hypothesis $H_{j \rightarrow i}^{(h)}$ of non-causality at horizon h from w_{jt} to w_{it} ($w_j \xrightarrow[h]{\rightarrow} w_i | I_{(j)}$), the asymptotic distribution of the corresponding statistic $\mathcal{W}[H_0(h)]$ is $\chi^2(p)$.

The problem now consists in estimating V_{ip} . Let $\hat{u}_i(h) = [\hat{u}_i^{(h)}(t+h) : t=0, \dots, T-h]'$ be the vector of OLS residuals from the regression (26), $\hat{g}_i^{(h)}(t+h) = W_p(t) \hat{u}_i^{(h)}(t+h)$, and set

$$R_i^{(h)}(\tau) = \frac{1}{T-h} \sum_{t=\tau}^{T-h} \hat{g}_i^{(h)}(t+h) \hat{g}_i^{(h)}(t+h-\tau)', \quad \tau = 0, 1, 2, \dots$$

If the innovations are i.i.d. or, more generally, if (22) holds, a natural estimator of V_{ip} , which would take into account the fact that the prediction errors $u^{(h)}(t+h)$ follow an MA($h-1$) process, is given by:

$$\hat{V}_{ip}^{(W)}(T) = R_i^{(h)}(0) + \sum_{\tau=1}^{h-1} [R_i^{(h)}(\tau) + R_i^{(h)}(\tau)'] .$$

Under regularity conditions studied by White (1999, Section 6.3),

$$\hat{V}_{ip}^{(W)}(T) - V_{ip} \xrightarrow[T \rightarrow \infty]{\mathbf{p}} 0.$$

A problem with $\hat{V}_{ip}^{(W)}(T)$ is that it is not necessarily positive-definite.

An alternative estimator which is automatically positive-semidefinite is the one suggested by

Doan and Litterman (1983), Gallant (1987) and Newey and West (1987):

$$\hat{V}_{ip}^{(NW)}(T) = R_i^{(h)}(0) + \sum_{\tau=1}^{m(T)-1} \kappa(\tau, m(T)) \left[R_i^{(h)}(\tau) + R_i^{(h)}(\tau)' \right], \quad (32)$$

where $\kappa(\tau, m) = 1 - [\tau / (m + 1)]$, $\lim_{T \rightarrow \infty} m(T) = \infty$, and $\lim_{T \rightarrow \infty} [m(T) / T^{1/4}] = 0$. Under the regularity conditions given by Newey and West (1987),

$$\hat{V}_{ip}^{(NW)}(T) - V_{ip} \xrightarrow{T \rightarrow \infty} 0.$$

Other estimators that could be used here includes various heteroskedasticity-autocorrelation-consistent (HAC) estimators; see Andrews (1991), Andrews and Monahan (1992), Cribari-Neto, Ferrari and Cordeiro (2000), Cushing and McGarvey (1999), Den Haan and Levin (1997), Hansen (1992), Newey and McFadden (1994), Wooldridge (1989).

The cost of having a simple procedure that sidestep all the nonlinearities associated with the non-causality hypothesis is a loss of efficiency. There are two places where we are not using all information. The constraints on the $\pi_k^{(h)}$'s are giving information on the ψ_j 's and we are not using it. We are also estimating the VAR by OLS and correcting the variance-covariance matrix instead of doing a GLS-type estimation. These two sources of inefficiencies could potentially be overcome but it would lead to less user-friendly procedures.

The asymptotic distribution provided by Proposition 2, may not be very reliable in finite samples, especially if we consider a VAR system with a large number of variables and/or lags. Due to autocorrelation, a larger horizon may also affect the size and power of the test. So an alternative to using the asymptotic distribution chi-square of $\mathcal{W}[H_0(h)]$, consists in using Monte Carlo test techniques [see Dufour (2005)] or bootstrap methods [see, for example, Paparoditis (1996), Paparoditis and Streitberg (1991), Kilian (1998a, 1998b)]. In view of the fact that the asymptotic distribution of $\mathcal{W}[H_0(h)]$ is nuisance-parameter-free, such methods yield asymptotically valid tests when applied to $\mathcal{W}[H_0(h)]$ and typically provide a much better control of test level in finite samples. It is also possible that using better estimates would improve size control, although this is not clear, for important size distortions can occur in multivariate regressions even when unbiased efficient estimators are available [see, for example, Dufour and Khalaf (2002)].

5. Causality tests based on nonstationary (p, h) -autoregressions

In this section, we study how the tests described in the previous section can be adjusted in order to allow for non-stationary possibly integrated processes. In particular, let us assume that

$$W(t) = \mu(t) + \eta(t), \quad (33)$$

$$\mu(t) = \delta_0 + \delta_1 t + \cdots + \delta_q t^q, \quad \eta(t) = \sum_{k=1}^p \pi_k \eta(t-k) + a(t), \quad (34)$$

$t = 1, \dots, T$, where $\delta_0, \delta_1, \dots, \delta_q$ are $m \times 1$ fixed vectors, and the process $\eta(t)$ is at most $I(d)$ where d is an integer greater than or equal to zero. Typical values for d are 0, 1 or 2. Note that these assumptions allow for the presence (or the absence) of cointegration relationships.

Under the above assumptions, we can also write:

$$W(t) = \gamma_0 + \gamma_1 t + \cdots + \gamma_q t^q + \sum_{k=1}^p \pi_k W(t-k) + a(t), \quad t = 1, \dots, T, \quad (35)$$

where $\gamma_0, \gamma_1, \dots, \gamma_q$ are $m \times 1$ fixed vectors (which depend on $\delta_0, \delta_1, \dots, \delta_q$, and π_1, \dots, π_p); see Toda and Yamamoto (1995). Under the specification (35), we have:

$$W(t+h) = \mu^{(h)}(t) + \sum_{k=1}^p \pi_k^{(h)} W(t+1-k) + u^{(h)}(t+h), \quad t = 0, \dots, T-h. \quad (36)$$

where $\mu^{(h)}(t) = \gamma_0^{(h)} + \gamma_1^{(h)} t + \cdots + \gamma_q^{(h)} t^q$ and $\gamma_0^{(h)}, \gamma_1^{(h)}, \dots, \gamma_q^{(h)}$ are $m \times 1$ fixed vectors. For $h = 1$, this equation is identical with (35). For $h \geq 2$, the errors $u^{(h)}(t+h)$ follow a $MA(h-1)$ process as opposed to being i.i.d.. For any integer j , we have:

$$\begin{aligned} W(t+h) &= \mu^{(h)}(t) + \sum_{\substack{k=1 \\ k \neq j}}^p \pi_k^{(h)} [W(t+1-k) - W(t+1-j)] \\ &\quad + \left(\sum_{k=1}^p \pi_k^{(h)} \right) W(t+1-j) + u^{(h)}(t+h), \end{aligned} \quad (37)$$

$$\begin{aligned} W(t+h) - W(t+1-j) &= \mu^{(h)}(t) + \sum_{\substack{k=1 \\ k \neq j}}^p \pi_k^{(h)} [W(t+1-k) - W(t+1-j)] \\ &\quad - \left(I_m - \sum_{k=1}^p \pi_k^{(h)} \right) W(t+1-j) + u^{(h)}(t+h), \end{aligned} \quad (38)$$

for $t = 0, \dots, T - h$. The two latter expressions can be viewed as extensions to (p, h) -autoregressions of the representations used by Dolado and Lütkepohl (1996, pp. 372-373) for VAR(p) processes. Further, on taking $j = p + 1$ in (38), we see that

$$\begin{aligned} W(t+h) - W(t-p) &= \mu^{(h)}(t) + \sum_{k=1}^p A_k^{(h)} \Delta W(t+1-k) \\ &\quad + B_p^{(h)} W(t-p) + u^{(h)}(t+h) \end{aligned} \quad (39)$$

where $\Delta W(t) = W(t) - W(t-1)$, $A_k^{(h)} = \sum_{j=1}^k \pi_j^{(h)}$, and $B_k^{(h)} = A_k^{(h)} - I_m$. Equation (39) may be interpreted as an error-correction form at the horizon h , with base $W(t-p)$.

Let us now consider the extended autoregression

$$W(t+h) = \mu^{(h)}(t) + \sum_{k=1}^p \pi_k^{(h)} W(t+1-k) + \sum_{k=p+1}^{p+d} \pi_k^{(h)} W(t+1-k) + u^{(h)}(t+h), \quad (40)$$

$t = d, \dots, T - h$. Under model (35), the actual values of the coefficient matrices $\pi_{p+1}^{(h)}, \dots, \pi_{p+d}^{(h)}$ are equal to zero ($\pi_{p+1}^{(h)} = \dots = \pi_{p+d}^{(h)} = 0$), but we shall estimate the (p, h) -autoregressions without imposing any restriction on $\pi_{p+1}^{(h)}, \dots, \pi_{p+d}^{(h)}$.

Now, suppose the process $\eta(t)$ is either $I(0)$ or $I(1)$, and we take $d = 1$ in (40). Then, on replacing p by $p + 1$ and setting $j = p$ in the representation (38), we see that

$$\begin{aligned} W(t+h) - W(t-p-1) &= \mu^{(h)}(t) + \sum_{k=1}^p \pi_k^{(h)} [W(t+1-k) - W(t-p-1)] \\ &\quad - B_{p+1}^{(h)} W(t-p-1) + u^{(h)}(t+h), \end{aligned} \quad (41)$$

where $B_{p+1}^{(h)} = (I_m - \sum_{k=1}^{p+1} \pi_k^{(h)})$. In the latter equation, $\pi_1^{(h)}, \dots, \pi_p^{(h)}$ all affect trend-stationary variables (in an equation where a trend is included along with the other coefficients). Using arguments similar to those of Sims et al. (1990), Park and Phillips (1989) and Dolado and Lütkepohl (1996), it follows that the estimates of $\pi_1^{(h)}, \dots, \pi_p^{(h)}$ based on estimating (41) by ordinary least squares (without restricting $B_{p+1}^{(h)}$) – or, equivalently, those obtained from (40) without restricting $\pi_{p+1}^{(h)}$ – are asymptotically normal with the same asymptotic covariance matrix as the one obtained for a stationary process of the type studied in section 4.² Consequently, the asymptotic distribution of the statistic $\mathcal{W}[H_{j \rightarrow i}^{(h)}]$ for testing the null hypothesis $H_{j \rightarrow i}^{(h)}$ of non-

²For related results, see also Choi (1993), Toda and Yamamoto (1995), Yamamoto (1996), Yamada and Toda (1998), and Kurozumi and Yamamoto (2000).

causality at horizon h from w_j to w_i ($w_j \xrightarrow{h} w_i | I_{(j)}$), based on estimating (40), is $\chi^2(p)$. When computing $H_{j \rightarrow i}^{(h)}$ as defined in (28), it is important that only the coefficients of $\pi_1^{(h)}, \dots, \pi_p^{(h)}$ are restricted (but not $\pi_{p+1}^{(h)}$).

If the process $\eta(t)$ is integrated up to order d , where $d \geq 0$, we can proceed similarly and add d extra lags to the VAR process studied. Again, the null hypothesis is tested by considering the restrictions entailed on $\pi_1^{(h)}, \dots, \pi_p^{(h)}$. Further, in view of the fact the test statistics are asymptotically pivotal under the null hypothesis, it is straightforward to apply bootstrap methods to such statistics. Note finally that the precision of the VAR estimates in such augmented regressions may eventually be improved with respect to the OLS estimates considered here by applying bias corrections such as those proposed by Kurozumi and Yamamoto (2000)]. Adapting and applying such corrections to (p, h) -autoregressions would go beyond the scope of the present paper.

6. Empirical illustration

In this section, we present an application of these causality tests at various horizons to macroeconomic time series. The data set considered is the one used by Bernanke and Mihov (1998) in order to study United States monetary policy. The data set considered consists of monthly observations on nonborrowed reserves (NBR , also denoted w_1), the federal funds rate (r , w_2), the GDP deflator (P , w_3) and real GDP (GDP , w_4). The monthly data on GDP and GDP deflator were constructed by state space methods from quarterly observations [see Bernanke and Mihov (1998) for more details]. The sample goes from January 1965 to December 1996 for a total of 384 observations. In what follows, all the variables were first transformed by a logarithmic transformation.

Before performing the causality tests, we must specify the order of the VAR model. First, in order to get apparently stationary time series, all variables were transformed by taking first differences of their logarithms. In particular, for the federal funds rate, this helped to mitigate the effects of a possible break in the series in the years 1979-1981.³ Starting with 30 lags, we then tested the hypothesis of K lags versus $K + 1$ lags using the LR test presented in Tiao and Box (1981). This led to a VAR(16) model. Tests of a VAR(16) against a VAR(K) for $K =$

³Bernanke and Mihov (1998) performed tests for arbitrary break points, as in Andrews (1993), and did not find significant evidence of a break point. They considered a VAR(13) with two additional variables (total bank reserves and Dow-Jones index of spot commodity prices and they normalize both reserves by a 36-month moving average of total reserves.)

17, . . . , 30 also failed to reject the VAR(16) specification, and the AIC information criterion [see McQuarrie and Tsai (1998, chapter 5)] is minimized as well by this choice. Calculations were performed using the Ox program (version 3.00) working on Linux [see Doornik (1999)].

Vector autoregressions of order p at horizon h were estimated as described in section 4 and the matrix $\hat{V}_{ip}^{(NW)}$, required to obtain covariance matrices, were computed using formula (32) with $m(T) - 1 = h - 1$.⁴ On looking at the values of the test statistics and their corresponding p -values at various horizons it quickly becomes evident that the $\chi^2(q)$ asymptotic approximation of the statistic \mathcal{W} in equation (31) is very poor. As a simple Monte Carlo experiment, we replaced the data by a 383×4 matrix of random draw from an $N(0, 1)$, ran the same tests and looked at the rejection frequencies over 1000 replications using the asymptotic critical value. The results are in Table 1a. We see important size distortions even for the tests at horizon 1 where there is no moving average part.

We next illustrate that the quality of the asymptotic approximation is even worse when we move away from an i.i.d. Gaussian setup to a more realistic case. We now take as the DGP the VAR(16) estimated with our data in first difference but we impose that some coefficients are zero such that the federal funds rate does not cause GDP up to horizon h and then we test the $r \xrightarrow[h]{} GDP$ hypothesis. The constraints of non-causality from j to i up to horizon h that we impose are:

$$\hat{\pi}_{ijl} = 0 \quad \text{for } 1 \leq l \leq p, \quad (42)$$

$$\hat{\pi}_{ikl} = 0 \quad \text{for } 1 \leq l \leq h, 1 \leq k \leq m. \quad (43)$$

Rejection frequencies for this case are given in Table 1b.

In light of these results we computed the p -values by doing a parametric bootstrap, *i.e.* doing an asymptotic Monte Carlo test based on a consistent point estimate [see Dufour (2005)]. The procedure to test the hypothesis $w_j \xrightarrow[h]{} w_i \mid I_{(j)}$ is the following.

1. An unrestricted VAR(p) model is fitted for the horizon one, yielding the estimates $\hat{\Pi}^{(1)}$ and $\hat{\Omega}$ for $\Pi^{(1)}$ and Ω .
2. An unrestricted (p, h) -autoregression is fitted by least squares, yielding the estimate $\hat{\Pi}^{(h)}$ of $\Pi^{(h)}$.

⁴The covariance estimator used here is relatively simple and exploits the truncation property (22). In view of the vast literature on HAC estimators [see Den Haan and Levin (1997) and Cushing and McGarvey (1999)], several alternative estimators for V_{ip} could have been considered (possibly allowing for alternative assumptions on the innovation distribution). It would certainly be of interest to compare the performances of alternative covariance estimators, but this would lead to a lengthy study, beyond the scope of the present paper.

Table 1. Rejection frequencies using the asymptotic distribution and the bootstrap procedure
a) i.i.d. Gaussian sequence

$h =$	1	2	3	4	5	6	7	8	9	10	11	12
Asymptotic												
5% level	27.0	27.8	32.4	36.1	35.7	42.6	47.9	48.5	51.0	55.7	59.7	63.6
10% level	37.4	39.4	42.2	46.5	47.8	52.0	58.1	59.3	60.3	66.3	69.2	72.5
Bootstrap												
5% level	5.5	5.7	4.7	6.5	4.0	5.1	5.5	3.9	4.7	6.1	5.2	3.8
10% level	10.0	9.1	10.1	10.9	9.6	10.6	10.2	9.4	9.5	10.9	10.3	8.9

b) VAR(16) without causality up to horizon h

Asymptotic												
5% level	24.1	27.9	35.8	37.5	55.9	44.3	52.3	55.9	54.1	60.1	62.6	72.0
10% level	35.5	38.3	46.6	47.2	65.1	55.0	64.7	64.6	64.8	69.8	72.0	79.0
Bootstrap												
5% level	6.0	5.1	3.8	6.1	4.6	4.7	4.4	4.5	4.3	6.3	4.9	5.8
10% level	9.8	8.8	8.7	10.4	10.3	9.9	8.7	7.4	10.3	11.1	9.3	9.7

3. The test statistic \mathcal{W} for testing noncausality at the horizon h from w_j to w_i [$H_{j \rightarrow i}^{(h)} : w_j \xrightarrow{h} w_i \mid I_{(j)}$] is computed. We denote by $\mathcal{W}_{j \rightarrow i}^{(h)}(0)$ the test statistic based on the actual data.
4. N simulated samples from (8) are drawn by Monte Carlo methods, using $\Pi^{(h)} = \hat{\Pi}^{(h)}$ and $\Omega = \hat{\Omega}$ [and the hypothesis that $a(t)$ is Gaussian]. We impose the constraints of non-causality, $\hat{\pi}_{ijk}^{(h)} = 0$, $k = 1, \dots, p$. Estimates of the impulse response coefficients are obtained from $\hat{\Pi}^{(1)}$ through the relations described in equation (5). We denote by $\mathcal{W}_{j \rightarrow i}^{(h)}(n)$ the test statistic for $H_{j \rightarrow i}^{(h)}$ based on the n -th simulated sample ($1 \leq n \leq N$).
5. The simulated p -value $\hat{p}_N[\mathcal{W}_{j \rightarrow i}^{(h)}(0)]$ is obtained, where

$$\hat{p}_N[x] = \left\{ 1 + \sum_{n=1}^N I[\mathcal{W}_{j \rightarrow i}^{(h)}(n) - x] \right\} / (N + 1),$$

$I[z] = 1$ if $z \geq 0$ and $I[z] = 0$ if $z < 0$.

6. The null hypothesis $H_{j \rightarrow i}^{(h)}$ is rejected at level α if $\hat{p}_N[\mathcal{W}_{j \rightarrow i}^{(h)}(0)] \leq \alpha$.

From looking at the results in Table 1, we see that we get a much better size control by using this bootstrap procedure. The rejection frequencies over 1000 replications (with $N = 999$) are very close to the nominal size. Although the coefficients ψ_j 's are functions of the π_i 's we do not constrain them in the bootstrap procedure because there is no direct mapping from $\pi_k^{(h)}$ to π_k and ψ_j . This certainly produces a power loss but the procedure remains valid because the $\hat{\psi}_j$'s are computed with the $\hat{\pi}_k$, which are consistent estimates of the true π_k both under the null and alternative hypothesis. To illustrate that our procedure has power for detecting departure from the null hypothesis of non-causality at a given horizon we ran the following Monte Carlo experiment. We again took a VAR(16) fitted on our data in first differences and we imposed the constraints (42) - (43) so that there was no causality from r to GDP up to horizon 12 (DGP under the null hypothesis). Next the value of one coefficient previously set to zero was changed to induce causality from r to GDP at horizons 4 and higher: $\pi_3(1, 3) = \theta$. As θ increases from zero to one the strength of the causality from r to GDP is higher. Under this setup, we could compute the power of our simulated test procedure to reject the null hypothesis of non-causality at a given horizon. In Figure 1, the power curves are plotted as a function of θ for the various horizons. The level of the tests was controlled through the bootstrap procedure. In this experiment we took again $N = 999$ and we did 1000 simulations. As expected, the power curves are flat at around 5% for horizons one to three since the null is true for these horizons. For horizons four and up we get the expected result that power goes up as θ moves from zero to one, and the power curves gets flatter as we increase the horizon.

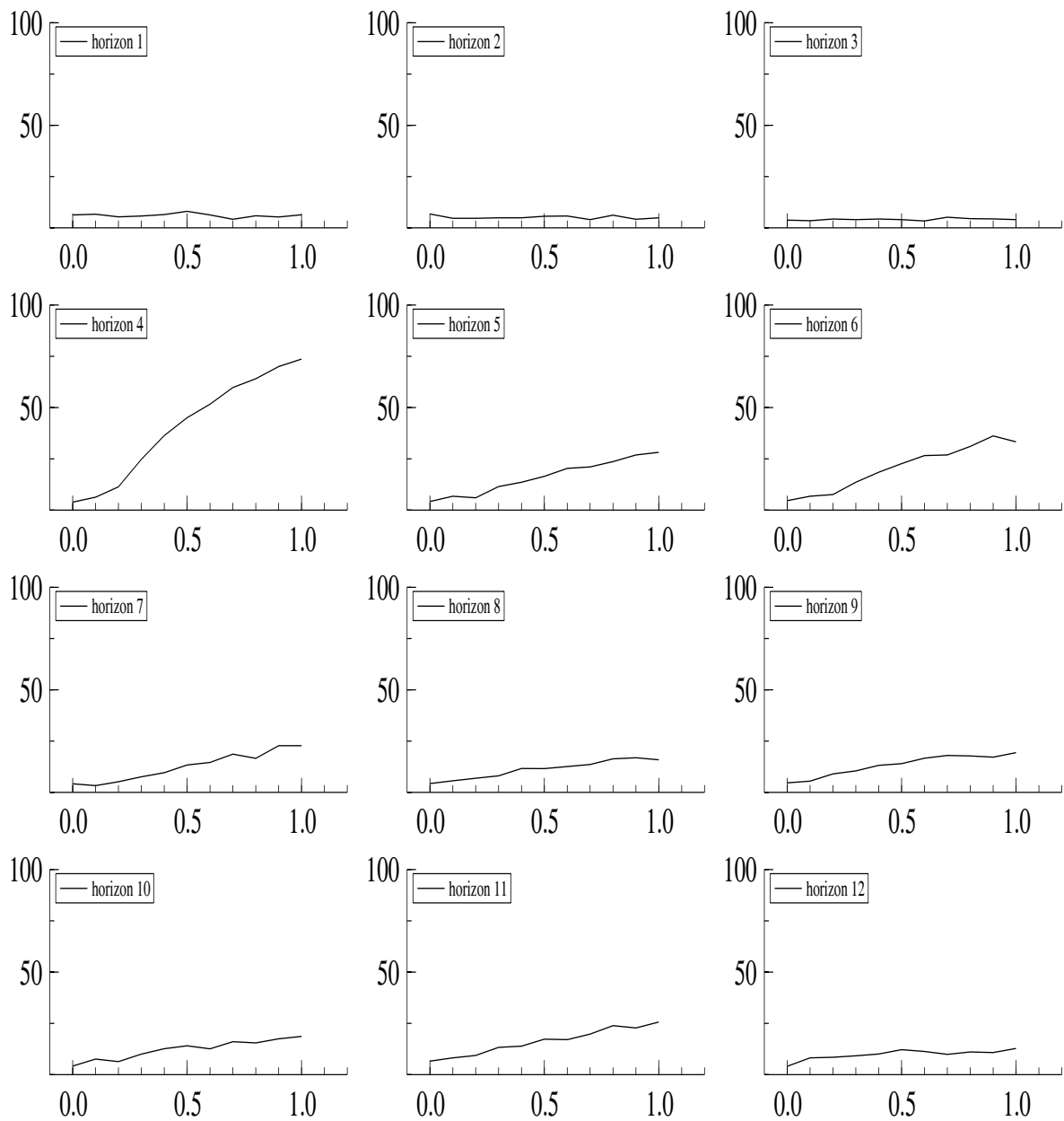


Figure 1. Power of the test at the 5% level for given horizons.
The abscissa (x axis) represents the values of θ .

Table 2. Causality tests and simulated p -values for series in first differences (of logarithm) for the horizons 1 to 12. p -values are reported in parenthesis.

h	\rightarrow	1	2	3	4	5	6	7	8	9	10	11	12
NBR	$\rightarrow r$	38.5205 (0.041)	26.3851 (0.240)	24.3672 (0.327)	22.5684 (0.395)	24.2294 (0.379)	27.0748 (0.313)	21.4347 (0.550)	17.2164 (0.799)	21.8217 (0.603)	18.4775 (0.780)	18.5379 (0.824)	19.6482 (0.789)
	$r \rightarrow$	22.5390 (0.386)	20.8621 (0.467)	17.2357 (0.706)	17.4222 (0.738)	17.8944 (0.734)	18.6462 (0.735)	25.2059 (0.495)	40.9896 (0.115)	41.6882 (0.142)	38.1656 (0.214)	41.2203 (0.206)	36.1278 (0.346)
NBR	$\rightarrow P$	50.5547 (0.004)	45.2498 (0.014)	49.1408 (0.009)	33.8545 (0.121)	33.8943 (0.162)	35.3923 (0.157)	39.1767 (0.099)	40.0218 (0.141)	36.4089 (0.234)	39.0916 (0.202)	28.4933 (0.485)	26.5439 (0.609)
	$P \rightarrow$	17.0696 (0.689)	18.9504 (0.601)	16.6880 (0.741)	21.2381 (0.550)	28.7264 (0.307)	20.0359 (0.664)	18.2291 (0.771)	23.6704 (0.596)	23.3419 (0.649)	27.8427 (0.492)	30.9171 (0.462)	36.5159 (0.357)
NBR	$\rightarrow GDP$	27.8029 (0.184)	25.0122 (0.302)	25.5123 (0.294)	23.6799 (0.393)	15.0040 (0.830)	17.5748 (0.757)	16.6781 (0.790)	19.6643 (0.746)	29.9020 (0.368)	34.0930 (0.300)	32.3917 (0.370)	34.8183 (0.357)
GDP	$\rightarrow NBR$	17.6338 (0.644)	20.6568 (0.520)	24.5334 (0.348)	18.3220 (0.670)	18.3123 (0.682)	32.8746 (0.211)	33.9979 (0.242)	39.6701 (0.145)	40.7356 (0.161)	26.3424 (0.551)	40.2262 (0.216)	53.3482 (0.092)
	$r \rightarrow$	32.2481 (0.104)	32.9207 (0.108)	32.0362 (0.138)	25.0124 (0.342)	25.2441 (0.383)	25.8110 (0.394)	29.2553 (0.328)	29.6021 (0.351)	36.0771 (0.241)	43.2116 (0.138)	30.2530 (0.422)	20.8982 (0.763)
	$P \rightarrow$	22.4385 (0.413)	16.4455 (0.670)	14.3073 (0.790)	14.2932 (0.826)	14.0148 (0.844)	16.7138 (0.774)	11.5599 (0.951)	16.5731 (0.809)	13.0697 (0.936)	14.5759 (0.909)	15.0317 (0.899)	24.4077 (0.659)
	$r \rightarrow$	26.3362 (0.262)	28.6645 (0.159)	35.9768 (0.073)	38.8680 (0.059)	37.6788 (0.082)	39.8145 (0.079)	64.1500 (0.006)	80.2396 (0.002)	89.6998 (0.002)	101.4143 (0.002)	105.9138 (0.001)	110.3551 (0.003)
GDP	$\rightarrow r$	42.5327 (0.016)	43.0078 (0.017)	49.8366 (0.005)	41.1082 (0.033)	41.7945 (0.044)	36.0965 (0.109)	29.5472 (0.263)	25.7898 (0.429)	27.9807 (0.366)	39.6366 (0.148)	39.0477 (0.178)	40.1545 (0.205)
	$P \rightarrow$	20.6903 (0.495)	24.1099 (0.322)	27.4106 (0.233)	23.3585 (0.423)	22.9095 (0.497)	18.5543 (0.713)	20.8172 (0.639)	23.6942 (0.555)	30.5340 (0.375)	28.8286 (0.414)	24.9477 (0.612)	25.7552 (0.629)
GDP	$\rightarrow P$	24.3368 (0.329)	24.4925 (0.365)	24.6125 (0.380)	22.8160 (0.470)	26.7900 (0.348)	40.0825 (0.081)	36.4855 (0.157)	49.6161 (0.058)	46.2574 (0.072)	36.3197 (0.262)	26.8520 (0.540)	24.0113 (0.666)

Table 3. Causality tests and simulated p -values for series in first differences (of logarithm) for the horizons 13 to 24

h	13	14	15	16	17	18	19	20	21	22	23	24
$NBR \rightarrow r$	20.7605 (0.783)	21.1869 (0.771)	18.6062 (0.852)	17.6750 (0.905)	22.2838 (0.811)	34.0098 (0.532)	28.7769 (0.696)	33.7855 (0.632)	66.1538 (0.143)	38.7272 (0.558)	28.8194 (0.803)	34.8015 (0.708)
$r \rightarrow NBR$	42.2924 (0.245)	43.8196 (0.254)	32.8113 (0.533)	28.1775 (0.707)	31.2326 (0.625)	31.7587 (0.651)	25.6942 (0.832)	34.6680 (0.682)	37.7470 (0.625)	51.4196 (0.406)	33.2897 (0.755)	27.3362 (0.869)
$NBR \rightarrow P$	38.0451 (0.336)	39.0293 (0.351)	43.2101 (0.315)	37.9049 (0.413)	22.4428 (0.835)	20.0337 (0.883)	38.8919 (0.488)	66.1541 (0.149)	55.0460 (0.255)	64.8568 (0.180)	54.8929 (0.332)	42.2509 (0.562)
$P \rightarrow NBR$	29.5457 (0.565)	29.4543 (0.667)	27.2425 (0.688)	28.7547 (0.703)	31.4851 (0.627)	44.1254 (0.400)	36.0319 (0.581)	47.1608 (0.427)	50.2743 (0.379)	45.8078 (0.486)	54.1961 (0.401)	56.7865 (0.399)
$NBR \rightarrow GDP$	30.3906 (0.493)	25.2694 (0.716)	29.7274 (0.603)	23.3347 (0.814)	16.7784 (0.942)	18.7922 (0.901)	26.8942 (0.752)	30.2550 (0.745)	39.0652 (0.545)	43.6661 (0.505)	56.9724 (0.314)	67.7764 (0.251)
$GDP \rightarrow NBR$	44.2264 (0.219)	41.5795 (0.292)	59.3301 (0.106)	65.7647 (0.086)	53.3579 (0.225)	50.4645 (0.282)	51.5163 (0.311)	43.1132 (0.466)	40.7051 (0.513)	36.3960 (0.659)	38.7162 (0.665)	40.8466 (0.644)
$r \rightarrow P$	27.3588 (0.588)	37.7170 (0.341)	37.2499 (0.381)	37.1005 (0.443)	27.3052 (0.700)	35.4128 (0.534)	39.2469 (0.477)	53.2675 (0.266)	57.1530 (0.229)	54.6753 (0.314)	69.6377 (0.164)	59.2184 (0.293)
$P \rightarrow r$	22.3720 (0.750)	26.3588 (0.608)	28.4930 (0.633)	23.0152 (0.794)	26.4211 (0.747)	36.3632 (0.512)	34.3922 (0.536)	17.4269 (0.956)	16.9652 (0.958)	31.9336 (0.715)	28.8377 (0.801)	30.0013 (0.795)
$r \rightarrow GDP$	123.8280 (0.001)	80.3638 (0.018)	95.0727 (0.005)	83.2773 (0.018)	76.6169 (0.033)	84.1068 (0.040)	76.8979 (0.055)	81.3505 (0.073)	65.5025 (0.161)	71.3334 (0.153)	63.1942 (0.241)	64.8020 (0.286)
$GDP \rightarrow r$	35.7086 (0.319)	33.0200 (0.420)	40.1713 (0.273)	30.3227 (0.545)	20.5582 (0.838)	19.1144 (0.898)	17.7386 (0.930)	22.6410 (0.856)	38.8175 (0.515)	38.7110 (0.565)	38.9987 (0.549)	25.0577 (0.860)
$P \rightarrow GDP$	9.4290 (0.988)	9.9870 (0.994)	12.1148 (0.974)	15.1682 (0.947)	14.2883 (0.970)	21.2701 (0.868)	29.0528 (0.708)	47.5841 (0.363)	66.5988 (0.162)	59.2137 (0.256)	71.5165 (0.165)	67.1851 (0.231)
$GDP \rightarrow P$	29.7642 (0.521)	37.0095 (0.351)	33.4676 (0.470)	42.2190 (0.328)	31.1573 (0.605)	52.3757 (0.238)	51.9567 (0.226)	40.3790 (0.459)	31.6598 (0.675)	54.0684 (0.244)	65.8284 (0.215)	53.0823 (0.342)

Table 4. Summary of causality relations at various horizons for series in first difference

h		1	2	3	4	5	6	7	8	9	10	11	12
<i>NBR</i>	\rightarrow	**											
<i>r</i>	\rightarrow	**											
<i>NBR</i>	\rightarrow	**	**	**				*					
<i>P</i>	\rightarrow	**											
<i>NBR</i>	\rightarrow	**											
<i>GDP</i>	\rightarrow	**											
<i>r</i>	\rightarrow	**											
<i>P</i>	\rightarrow	**											
<i>r</i>	\rightarrow			*	*	*	*	**	**	**	**	**	**
<i>GDP</i>	\rightarrow	**	**	**	**	**							
<i>P</i>	\rightarrow	**											
<i>GDP</i>	\rightarrow						*		*	*			
h		13	14	15	16	17	18	19	20	21	22	23	24
<i>NBR</i>	\rightarrow	**											
<i>r</i>	\rightarrow	**											
<i>NBR</i>	\rightarrow	**											
<i>P</i>	\rightarrow	**											
<i>NBR</i>	\rightarrow	**											
<i>GDP</i>	\rightarrow	**											
<i>r</i>	\rightarrow	**											
<i>P</i>	\rightarrow	**											
<i>r</i>	\rightarrow	**	**	**	**	**	**	*	*				
<i>GDP</i>	\rightarrow	**											
<i>P</i>	\rightarrow	**											
<i>GDP</i>	\rightarrow	**											

Note _ The symbols * and ** indicate rejection of the non-causality hypothesis at the 10% and 5% levels respectively.

Now that we have shown that our procedure does have power we present causality tests at horizon one to 24 for every pair of variables in tables 2 and 3. For every horizon we have twelve causality tests and we group them by pairs. When we say that a given variable cause or does not cause another, it should be understood that we mean the growth rate of the variables. The p -values are computed by taking $N = 999$. Table 4 summarize the results by presenting the significant results at the 5% and 10% levels.

The first thing to notice is that we have significant causality results at short horizons for some pairs of variables while we have it at longer horizons for other pairs. This is an interesting illustration of the concept of causality at horizon h of Dufour and Renault (1998).

The instrument of the central bank, the nonborrowed reserves, cause the federal funds rate at horizon one, the prices at horizon 1, 2, 3 and 9 (10% level). It does not cause the other two variables at any horizon and except the GDP at horizon 12 and 16 (10% level) nothing is causing it. We see that the impact of variations in the nonborrowed reserves is over a very short term. Another variable, the GDP, is also causing the federal funds rates over short horizons (one to five months).

An interesting result is the causality from the federal funds rate to the GDP. Over the first few months the funds rate does not cause GDP, but from horizon 3 (up to 20) we do find significant causality. This result can easily be explained by, e.g. the theory of investment. Notice that we have the following indirect causality. Nonborrowed reserves do not cause GDP directly over any horizon, but they cause the federal funds rate which in turn causes GDP. Concerning the observation that there are very few causality results for long horizons, this may reflect the fact that, for stationary processes, the coefficients of prediction formulas converge to zero as the forecast horizon increases.

Using the results of Proposition 4.5 in Dufour and Renault (1998), we know that for this example the highest horizon that we have to consider is 33 since we have a VAR(16) with four time series. Causality tests for the horizons 25 through 33 were also computed but are not reported. Some p -values smaller or equal to 10% are scattered over horizons 30 to 33 but no discernible pattern emerges.

We next consider extended autoregressions to illustrate the results of section 5. To cover the possibility that the first difference of the logarithm of the four series may not be stationary, we ran extended autoregressions on the series analyzed. Since we used a VAR(16) with non-zero mean for the first difference of the series a VAR(17), *i.e.* $d = 1$, with a non-zero mean was fitted. The Monte Carlo samples with $N = 999$ are drawn in the same way as before except

that the constraints on the VAR parameters at horizon h is $\hat{\pi}_{jik}^{(h)} = 0$ for $k = 1, \dots, p$ and not $k = 1, \dots, p + d$.

Results of the extended autoregressions are presented in Table 5 (horizons 1 to 12) and 6 (horizons 13 to 24). Table 7 summarize these results by presenting the significant results at the 5% and 10% level. These results are very similar to the previous ones over all the horizons and variable every pairs. A few causality tests are not significant anymore ($GDP \rightarrow r$ at horizon 5, $r \rightarrow GDP$ at horizons 5 and 6) and some causality relations are now significant ($r \rightarrow P$ at horizon one) but we broadly have the same causality patterns.

7. Conclusion

In this paper, we have proposed a simple linear approach to the problem of testing non-causality hypotheses at various horizons in finite-order vector autoregressive models. The methods described allow for both stationary (or trend-stationary) processes and possibly integrated processes (which may involve unspecified cointegrating relationships), as long as an upper bound is set on the order of integration. Further, we have shown that these can be easily implemented in the context of a four-variable macroeconomic model of the U.S. economy.

Several issues and extensions of interest warrant further study. The methods we have proposed were, on purpose, designed to be relatively simple to implement. This may, of course, involve efficiency losses and leave room for improvement. For example, it seems quite plausible that more efficient tests may be obtained by testing directly the nonlinear causality conditions described in Dufour and Renault (1998) from the parameter estimates of the VAR model. However, such procedures will involve difficult distributional problems and may not be as user-friendly as the procedures described here. Similarly, in nonstationary time series, information about integration order and the cointegrating relationships may yield more powerful procedures, although at the cost of complexity. These issues are the topics of on-going research.

Another limitation comes from the fact we consider VAR models with a known finite order. We should however note that the asymptotic distributional results established in this paper continue to hold as long as the order p of the model is selected according to a consistent order selection rule [see Dufour, Ghysels and Hall (1994), Pötscher (1991)]. So this is not an important restriction. Other problems of interest would consist in deriving similar tests applicable in the context of VARMA or VARIMA models, as well as more general infinite-order vector autoregressive models, using finite-order VAR approximations based on data-dependent truncation

Table 5. Causality tests and simulated p -values for extended autoregressions at the horizons 1 to 12

h	1	2	3	4	5	6	7	8	9	10	11	12
$NBR \rightarrow r$	37.4523 (0.051)	24.0148 (0.337)	24.9365 (0.309)	22.9316 (0.410)	25.4508 (0.333)	26.6763 (0.333)	19.5377 (0.660)	19.3805 (0.688)	21.9278 (0.609)	19.8541 (0.716)	21.2947 (0.710)	19.5569 (0.795)
$r \rightarrow NBR$	25.9185 (0.281)	17.7977 (0.650)	16.6185 (0.747)	17.3820 (0.759)	19.6425 (0.673)	20.2032 (0.676)	39.8496 (0.129)	44.0172 (0.075)	43.9928 (0.130)	39.4217 (0.207)	36.7802 (0.286)	35.0883 (0.375)
$NBR \rightarrow P$	50.8648 (0.004)	44.8472 (0.015)	56.5028 (0.007)	33.8112 (0.152)	36.1026 (0.124)	42.1714 (0.061)	38.7204 (0.152)	38.8076 (0.161)	35.3353 (0.259)	33.8049 (0.316)	28.6205 (0.498)	26.9863 (0.609)
$P \rightarrow NBR$	20.7491 (0.506)	17.2772 (0.704)	16.3891 (0.764)	26.1610 (0.339)	29.7329 (0.276)	18.7583 (0.730)	22.7821 (0.605)	23.1743 (0.629)	27.0076 (0.508)	24.3763 (0.655)	31.1869 (0.484)	38.3349 (0.336)
$NBR \rightarrow GDP$	27.2102 (0.224)	23.8072 (0.346)	25.9561 (0.317)	24.3384 (0.402)	14.4893 (0.837)	17.8928 (0.732)	15.9036 (0.821)	17.9592 (0.817)	30.7472 (0.368)	33.5333 (0.333)	33.7687 (0.355)	36.3343 (0.348)
$GDP \rightarrow NBR$	16.1322 (0.746)	19.4471 (0.571)	20.6798 (0.537)	19.1035 (0.658)	29.1229 (0.269)	36.1053 (0.166)	37.1194 (0.167)	40.3578 (0.163)	43.7835 (0.138)	37.7337 (0.247)	48.3004 (0.125)	52.2442 (0.123)
$r \rightarrow P$	32.5147 (0.100)	31.2909 (0.128)	25.6717 (0.352)	24.4449 (0.390)	21.7640 (0.568)	25.4747 (0.397)	27.6313 (0.353)	31.4581 (0.311)	43.2611 (0.111)	38.2020 (0.212)	30.6394 (0.420)	19.9687 (0.812)
$P \rightarrow r$	22.7374 (0.415)	15.9453 (0.704)	15.2001 (0.762)	15.1933 (0.790)	16.2334 (0.768)	15.7472 (0.830)	13.4196 (0.905)	15.2506 (0.859)	14.4324 (0.909)	15.8589 (0.887)	21.3637 (0.749)	21.9949 (0.731)
$r \rightarrow GDP$	27.0435 (0.244)	29.5913 (0.158)	37.5271 (0.061)	35.7130 (0.094)	34.9901 (0.117)	35.1715 (0.164)	79.9402 (0.002)	92.6009 (0.001)	94.9068 (0.004)	107.6638 (0.001)	108.0581 (0.003)	138.0570 (0.001)
$GDP \rightarrow r$	41.8475 (0.032)	41.9449 (0.019)	44.7597 (0.019)	38.0358 (0.060)	30.4776 (0.209)	28.1840 (0.286)	30.5867 (0.250)	27.0745 (0.369)	26.5876 (0.431)	39.9502 (0.157)	39.8618 (0.181)	33.4855 (0.347)
$P \rightarrow GDP$	23.7424 (0.368)	26.7148 (0.237)	24.6605 (0.342)	24.6507 (0.373)	23.3233 (0.479)	19.8483 (0.668)	20.3581 (0.666)	28.9318 (0.376)	29.0384 (0.427)	27.2509 (0.505)	26.3318 (0.558)	22.6991 (0.740)
$GDP \rightarrow P$	25.1264 (0.278)	24.1941 (0.358)	25.5683 (0.350)	19.2127 (0.639)	37.5984 (0.108)	38.0318 (0.110)	37.6254 (0.153)	45.2219 (0.088)	45.2458 (0.092)	36.9912 (0.237)	23.1687 (0.647)	22.9934 (0.698)

Table 6. Causality tests and simulated p -values for extended autoregressions at the horizons 13 to 24

h	13	14	15	16	17	18	19	20	21	22	23	24
$NBR \rightarrow r$	19.5845 (0.814)	30.2847 (0.513)	17.4893 (0.901)	24.9745 (0.750)	21.2814 (0.858)	27.4800 (0.706)	36.9567 (0.537)	27.8970 (0.770)	59.3731 (0.236)	34.9153 (0.640)	27.3241 (0.832)	31.9143 (0.771)
$r \rightarrow NBR$	41.7360 (0.282)	47.8501 (0.212)	30.9613 (0.592)	29.1809 (0.692)	35.7654 (0.540)	22.9966 (0.847)	29.0869 (0.757)	29.1396 (0.766)	38.0675 (0.665)	52.5512 (0.401)	29.6133 (0.838)	25.4872 (0.917)
$NBR \rightarrow P$	36.7787 (0.359)	37.7357 (0.366)	33.4273 (0.512)	35.3241 (0.455)	20.6330 (0.860)	23.6911 (0.825)	44.5735 (0.379)	55.5420 (0.249)	53.7340 (0.290)	68.0270 (0.165)	52.3332 (0.363)	47.5614 (0.481)
$P \rightarrow NBR$	29.5049 (0.582)	39.2076 (0.401)	18.0831 (0.920)	30.6486 (0.671)	39.5517 (0.441)	34.8363 (0.606)	39.9608 (0.535)	43.6563 (0.471)	40.6713 (0.551)	44.0254 (0.553)	61.6914 (0.296)	62.9346 (0.286)
$NBR \rightarrow GDP$	30.9525 (0.501)	22.0737 (0.822)	30.1165 (0.599)	22.7429 (0.793)	17.2546 (0.937)	22.2686 (0.863)	28.6752 (0.749)	31.8817 (0.717)	39.5031 (0.591)	53.6466 (0.367)	58.8413 (0.327)	79.0569 (0.153)
$GDP \rightarrow NBR$	46.4538 (0.186)	38.0424 (0.347)	64.2269 (0.071)	60.8792 (0.125)	57.5798 (0.169)	57.2237 (0.205)	42.0851 (0.453)	43.1683 (0.491)	43.4379 (0.501)	41.7141 (0.568)	39.8292 (0.631)	41.7123 (0.632)
$r \rightarrow P$	30.7883 (0.503)	37.3585 (0.364)	30.4331 (0.566)	35.8788 (0.448)	26.9960 (0.712)	39.2961 (0.429)	44.7334 (0.358)	46.6740 (0.352)	56.9680 (0.223)	53.1830 (0.310)	67.3818 (0.182)	61.2522 (0.241)
$P \rightarrow r$	25.3085 (0.654)	33.1027 (0.450)	28.7362 (0.624)	33.5758 (0.546)	33.8991 (0.525)	39.3031 (0.451)	29.4228 (0.707)	14.3095 (0.984)	17.8588 (0.957)	29.9572 (0.764)	25.0347 (0.883)	29.1903 (0.825)
$r \rightarrow GDP$	109.5052 (0.002)	84.9556 (0.011)	88.4433 (0.021)	80.9836 (0.016)	79.9549 (0.031)	75.1199 (0.073)	94.8986 (0.026)	65.5929 (0.147)	67.2913 (0.136)	71.6331 (0.164)	75.5123 (0.151)	67.4315 (0.254)
$GDP \rightarrow r$	34.8185 (0.340)	41.4218 (0.264)	38.2761 (0.318)	28.5326 (0.606)	22.6116 (0.809)	16.7992 (0.923)	20.8097 (0.866)	31.8769 (0.644)	38.7083 (0.519)	34.4663 (0.649)	35.6279 (0.637)	20.5007 (0.949)
$P \rightarrow GDP$	8.8039 (0.995)	8.9511 (0.995)	17.0182 (0.933)	9.7608 (0.995)	16.4772 (0.923)	19.6942 (0.916)	46.3240 (0.349)	41.7429 (0.484)	64.5928 (0.207)	60.2875 (0.250)	56.0315 (0.332)	72.2823 (0.215)
$GDP \rightarrow P$	23.8773 (0.709)	39.5163 (0.331)	34.3832 (0.468)	34.5734 (0.488)	36.8350 (0.472)	54.8088 (0.212)	54.8048 (0.218)	36.4102 (0.557)	29.4543 (0.739)	58.0095 (0.254)	58.1341 (0.265)	59.5283 (0.286)

Table 7. Summary of causality relations at various horizons for series in first difference with extended autoregressions

<i>h</i>		1	2	3	4	5	6	7	8	9	10	11	12
<i>NBR</i>	\nrightarrow <i>r</i>	*											
<i>r</i>	\nrightarrow <i>NBR</i>								*				
<i>NBR</i>	\nrightarrow <i>P</i>	**	**	**			*						
<i>P</i>	\nrightarrow <i>NBR</i>												
<i>NBR</i>	\nrightarrow <i>GDP</i>												
<i>GDP</i>	\nrightarrow <i>NBR</i>												
<i>r</i>	\nrightarrow <i>P</i>	*											
<i>P</i>	\nrightarrow <i>r</i>												
<i>r</i>	\nrightarrow <i>GDP</i>			*	*			**	**	**	**	**	**
<i>GDP</i>	\nrightarrow <i>r</i>	**	**	**	*								
<i>P</i>	\nrightarrow <i>GDP</i>												
<i>GDP</i>	\nrightarrow <i>P</i>								*	*			
<i>h</i>		13	14	15	16	17	18	19	20	21	22	23	24
<i>NBR</i>	\nrightarrow <i>r</i>												
<i>r</i>	\nrightarrow <i>NBR</i>												
<i>NBR</i>	\nrightarrow <i>P</i>												
<i>P</i>	\nrightarrow <i>NBR</i>												
<i>NBR</i>	\nrightarrow <i>GDP</i>												
<i>GDP</i>	\nrightarrow <i>NBR</i>			*									
<i>r</i>	\nrightarrow <i>P</i>												
<i>P</i>	\nrightarrow <i>r</i>												
<i>r</i>	\nrightarrow <i>GDP</i>	**	**	**	**	**	*	**					
<i>GDP</i>	\nrightarrow <i>r</i>												
<i>P</i>	\nrightarrow <i>GDP</i>												
<i>GDP</i>	\nrightarrow <i>P</i>												

Note _ The symbols * and ** indicate rejection of the non-causality hypothesis at the 10% and 5% levels respectively.

rules [such as those used by Lütkepohl and Poskitt (1996) and Lütkepohl and Saikkonen (1997)]. These problems are also the topics of on-going research.

Acknowledgements

This work was supported by the Canada Research Chair Program (Chair in Econometrics, Université de Montréal), the Alexander-von-Humboldt Foundation (Germany), the Institut de finance mathématique de Montréal (IFM2), the Canadian Network of Centres of Excellence [program on *Mathematics of Information Technology and Complex Systems* (MITACS)], the Canada Council for the Arts (Killam Fellowship), the Natural Sciences and Engineering Research Council of Canada, the Social Sciences and Humanities Research Council of Canada, the Fonds de recherche sur la société et la culture (Québec), and the Fonds de recherche sur la nature et les technologies (Québec). The authors thank Jörg Breitung, Helmut Lütkepohl, Peter Schotman, participants at the EC2 Meeting on Causality and Exogeneity in Louvain-la-Neuve (December 2001), two anonymous referees, and the Editors Luc Bauwens, Peter Boswijk and Jean-Pierre Urbain for several useful comments. This paper is a revised version of Dufour and Renault (1995).

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