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THE SEASONALITY OF THE ITALIAN COST OF LIVING INDEX

by Gianluca Cubadda[*] and Roberto Sabbatini[**]

Abstract

In the case of the consumer price index the importance of a seasonally adjusted measure lies in the possibility of promptly identifying turning points in inflation. The focus of this paper is the so-called "cost-of-living" index, the most used measure of Italian inflation. Our aim is to identify and test a model of its seasonal component, which is important in explaining the overall variability of the index. Most seasonality is actually spurious, being essentially linked to the fact that the data on a significant number of the items included in the cost-of-living index are collected quarterly. It is therefore very important to take this into account in order to adjust the cost-of-living index for seasonality; the type of seasonality involved must also be considered, since different methods need to be used depending on the characteristics of the series considered. On the basis of this preliminary analysis, the method implemented in the paper to adjust the series for seasonality is the model-based procedure known as SEATS. The measure recommended for monitoring the performance of inflation is the quarterly growth rate of the seasonally adjusted index, for two main reasons: first, a significant price change is always considered for all the series (i.e., for both monthly and quarterly series included in the cost-of-living index); second, it allows us to smooth out irregularities in price changes.

Contents

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1. Introduction[1]

From the point of view of the policy maker, the main aim of a statistical analysis of economic time series is to provide information which can facilitate the monitoring of the underlying trends in the economy. To this end, the use of seasonally adjusted measures is very helpful, as the short-term variability of macroeconomic time series is often strongly affected by seasonal factors. The concept of seasonality is intuitively clear, although no universally accepted definition exists (Hylleberg, 1992; Den Butter and Fase, 1991): seasonal variations are "regular" changes that occur at yearly intervals. In the case of the consumer price index, which is the focus of this paper, the importance of a seasonally adjusted measure lies in the possibility of promptly identifying the turning points of inflation, which cannot readily be identified either by changes over the previous period calculated on the raw data (which are affected by seasonal factors), or by changes over the corresponding period of the previous year (which reveal turning points long after they have occurred, since they are affected by developments over the whole year).

The focus of this paper is the so-called "cost-of-living" index (henceforth, CLI), the most frequently used measure of Italian inflation. Our aim is to identify and test a model of its seasonal component, so as to obtain a prompt signal of inflationary tendencies. The paper is divided into three sections. Section 2 discusses the sources of CLI seasonality and the main practical problems related to its

1 Helpful comments from Silvia Fabiani, Augustin Maravall, Christophe Planas, Paolo Sestito and an anonymous referee are gratefully acknowledged. We would also like to thank Daniela Falcone for research assistance. We can be contacted by e-mail at: cubadda@axrma.uniroma1.it; i8352157@interbusiness.it.

identification. Section 3 presents an analysis of the statistical characteristics of the seasonal component. Section 4 shows the empirical results of the seasonal adjustment procedure chosen and compares it with possible alternatives.

2. The seasonal nature of the cost-of-living index

Some idea of the significance of seasonal components can be obtained by regressing the monthly increases in the CLI[2] on 12 seasonal dummies; in the period from 1981 to 1995, seasonal dummies explain around 20 per cent of the total variability of monthly changes in the index. The importance of seasonality in explaining the short-term variability of the CLI has remained fairly stable over time, with the exception of the seventies (Figure 1).

Part of this seasonality is nonetheless spurious, in that it relates to prices of a significant share of items included in the CLI that are collected every three months

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[2] The CLI is a fixed-weight price index which refers only to the bundle of goods and services consumed by wage and salary earners. In principle, it might be worth investigating other indicators of Italian inflation as well: the consumer price index (CPI), which embraces the whole population, and the domestic households' consumption deflator (HCD), computed within the national accounts. In practice, however, several factors make the CLI, for the moment, a more influential index than either the CPI or the HCD. The CLI is available at the very beginning of the "following" month (with provisional data coming out around the 20th of the "current" month), whereas the CPI is usually only available with a delay of 1-2 months. Moreover, the CLI is available for the entire post-war period and, since it is used typically for administrative purposes (e.g., indexation of rents, disbursments for public works, alimony), it is the best known statistic in Italy. Finally, only the CLI is available net of changes in indirect taxes, a component that is best excluded when attempting to identify "core" inflationary impulses. In practice, both the CLI and the CPI show a very similar pattern. They might be expected to show a very similar seasonal pattern as well, being based on common methodologies and, at least partly, on the same set of raw data. As far as the HCD is concerned, although it provides the most accurate measure of inflation because of its variable weights and its complete representativeness, it is frequently revised and is available only quarterly, with a delay of several months.

(Table 1);[3] when prices are not collected, the latest available price is used. For practical purposes, "rents" (henceforth, R_t) are kept separate from "other" quarterly data (henceforth, Q_t), as the months in which data are collected differ in the two cases.[4] In the case of the official monthly Q_t and R_t series, monthly seasonal dummies explain 41 and 55 per cent respectively of their short-term variability in the period 1981-1995, compared with 15 per cent for the series which are "genuinely" observed each month (henceforth, M_t).[5] However, if (more correctly) Q_t and R_t are considered as purely quarterly series (as already mentioned, by construction each series has the same value within each quarter), the portion of seasonality captured by quarterly seasonal dummies is not very different from that of M_t (respectively, 13 and 26 per cent).

The following is a more formal presentation of how the existence of prices that are held approximately constant in each quarter may induce a spurious quarterly cycle in the growth rates of the CLI quarterly components. Let y_t represent the *growth rates* of a monthly series; let us then suppose that y_t is generated by the following process:

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[3] Items whose prices are collected every three months are mostly durables, a few market services and rents. "Medical tests", whose prices are collected every six months, have not been considered; their total weight is equal to 0.1 per cent in the 1992 basket (weights are rescaled to 100).

[4] Prices are collected in January, April, July and October in the case of rents, and in February, May, August and November in the case of "other" quarterly data.

[5] Disaggregated data for the CLI are only available to us from 1990. In order to consider a longer period of time, quarterly series and rents have been estimated for the years 1981-89 on the basis of the corresponding indices relative to the CPI; the monthly aggregate has been obtained as a residual (it is worth noting that in order to calculate these aggregates, varying base-year weights for the CLI have been used).

$$(1 - \rho L)\, y_t = \varepsilon_t$$

where L is the usual lag operator, ε_t is a white noise process and $\rho \in (0,1)$. Thus, y_t is a stationary AR(1) process with no seasonal behavior, given that its spectral density has a single peak at the zero frequency. However, the process y_t is not directly observable. The following series is observed instead:

$$z_t = y_t^* + v_t^*$$

where

$$v_t^* = \begin{cases} 0 & \text{if } t = 3[t/3]+1 \\ v_t & \text{otherwise} \end{cases}$$

v_t is a white noise process independent of ε_t, [.] denotes the integer part function, and

$$y_t^* = \begin{cases} y_t & \text{if } t = 3[t/3]+1 \\ 0 & \text{otherwise} \end{cases}$$

It is worth noting that monthly growth rates of the series Q_t and R_t are zero within the reference quarter, apart from occasional small approximation errors.

It is easy to verify that the second-order moments of series z_t are time-dependent. In fact, we have:

$$E(z_t\, z_{t-j}) = \begin{cases} \sigma_y^2 & \text{if } j = 0 \text{ and } t = 3\left[t / 3\right] + 1 \\ \sigma_v^2 & \text{if } j = 0 \text{ and } t \neq 3\left[t / 3\right] + 1 \\ \rho^j\, \sigma_y^2 & \text{if } j = 3\,n \text{ for } n \in N \text{ and } t = 3\left[t / 3\right] + 1 \\ 0 & \text{otherwise} \end{cases}$$

where σ_y^2 and σ_v^2 are, respectively, the variances of the processes y_t and v_t. Hence, the process z_t is not weakly stationary. However, if z_t were incorrectly assumed to be stationary, one would observe a sample autocorrelation function with values significantly different from zero only at lags $3j$ ($j = 1,2,\ldots$). Note that this is exactly the pattern shown, for instance, by the autocorrelation function of the growth rates of the Q_t series observed monthly (Figure 2).

As mentioned earlier, when all the series are considered at their actual frequency of observation, the seasonal factors (as captured by seasonal dummies) appear much less important. This result can be at least partly attributed to the following characteristics of the CLI. Firstly, seasonality of food prices is reduced because the Italian National Institute of Statistics (Istat) excludes items subject to extreme price changes from the computation of the fruit and vegetable indices in order to reduce volatility. Secondly, the high price volatility of energy products, which represent around 5 per cent of the monthly series, is typically more closely related to incidental factors than to seasonality.[6] Results reported in Table 2 (for the period 1990-95, since raw data for the CLI are available to us only from 1990) confirm that the contribution of seasonal dummies to explaining the short-term variability of the series

[6] Further confirmation of the small contribution of seasonality to explaining price changes for energy and fresh food products comes from the very poor performance (on the basis of the standard diagnostic) of the X11-ARIMA procedure applied to these aggregates.

considered is actually smaller for food and energy products, as well as most services, especially those whose prices are observed monthly.

The presence in the CLI basket of series with different frequencies of data-collection provides an "obvious" solution to the general issue of choosing between the so-called "direct" and "indirect" approaches (Bodo and Pellegrini, 1993). The former approach, which involves the direct seasonal adjustment of an aggregated series, has many advantages: it is easier to manage, it identifies the seasonal model more accurately because more sophisticated methods can be applied and, typically, it is possible to operate on a longer time series, which are normally available at the aggregate level. Moreover, where there are compensatory effects between seasonal patterns in the disaggregated series, a better estimate of the overall seasonal component can be derived. On the other hand, analysis of individual series ("indirect" approach) potentially leads to better identification of the seasonal models, thanks to information on the frequency of data sampling and, more in general, on the nature of the market for each type of good and service. The choice of one or other approach is essentially an empirical issue. In our case, a fair "compromise" between the advantages and disadvantages of the two solutions can be achieved by looking separately at the two quarterly series and at the monthly series.

An important implication of the above is that a significant price change for all the series results *only* when the quarterly rate of change of the seasonally adjusted components is considered.[7] Such a measure also has the important advantage of smoothing irregularities in the monthly

[7] Such a measure (which corresponds to the use of filter $(1-L^3)$) has been computed by first re-aggregating the three components considered (seasonally adjusted individually) with base-year varying weights.

price changes; in effect, an irregular price change in one month is often compensated by an opposite change in the following month (as typically happens with prices of fresh food products), without any change in the underlying inflationary tendencies. A "genuine" monthly signal on inflation is only provided by the seasonally adjusted M_t series; this is in fact significant, since the M_t series represents around 80 per cent of the CLI basket.

As regards the choice of time interval, the selection of a longer interval might facilitate statistical inference (e.g., because of the reduced influence of random events); however, the main disadvantages of this solution are both "genuine" structural changes over time in the seasonal pattern and the possible presence of "spurious" changes, typically due to methodological breaks in the original series (e.g., a relevant base year change or a change in the methodology underlying the computation of elementary indices). Both sources of change may "destabilize" the process of seasonal adjustment. We have selected the period 1981-1995, during which no major methodological changes occurred. The seasonal pattern might also change over time following the average level of inflation, and thus it is preferable to select a time interval during which inflation is fairly stable. Inflation has in fact been reasonably stable (around 5 per cent a year) only since 1986. However, the selection of too short a period might seriously affect the test procedure implemented in the next section, as well as the estimation of the seasonal pattern. In order to appreciate whether major changes occur in our results if a shorter period of time is considered, an empirical analysis was also carried out for the period 1986-1995 period. The results, which are not reported in the paper, are not significantly different from those presented in the two sections that follow.

Finally, a decision has to be made concerning the preliminary treatment of outliers, whose presence would distort the estimates of the various components. Where *ex ante* information is available, it is preferable to use it rather than implement an automatic procedure.[8] *Ex ante* information was used to treat changes in indirect taxes, which lead to one-off rises in prices, without, by themselves, changing the underlying inflationary trends. A preliminary adjustment for (such) outliers has been carried out using the procedure described in Gavosto, Sabbatini and Sestito (1996) and Banca d'Italia (1996). Briefly, in order to identify the seasonal component of the index, the irregular component attributable to tax changes is removed and reintroduced on completion of the seasonal adjustment procedure in the case of the CLI including tax changes. In order to take into account the likely delay in the full pass-through of tax changes, their effects are identified by considering only a fraction of the whole change in taxes. The empirical rule of thumb used to calculate a "modified" index removes 50, 30 and 20 per cent of VAT changes in the first, second and third months respectively; by contrast, excise duties are included in full, since their impact on prices is generally more immediate. The seasonal adjustment procedure is then applied to this modified index, adding back at the end the tax factors removed during the first stage. The reason for adopting this strategy is that tax changes no longer affect the identification of seasonal factors, since they are considered as a totally irregular component.

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[8] The major risk associated with the latter solution is that at first, until other data are available, an automatic outlier detection procedure might not be able to recognise the earlier figure as an actual outlier, leading to a spurious revision of the seasonal component. When more information became available, the outlier would be detected as such and the seasonal component further revised.

The same arguments hold for the CLI excluding changes in indirect taxes, which is particularly useful from the point of view of the policy maker, since such tax changes produce one-off variations in inflation that should not affect its underlying pattern. Istat compiles such an index on the assumption that changes in VAT rates and excise duties are passed through to prices immediately and in full. As noted previously, the pass-through of changes in VAT rates to prices probably takes place over several months. For instance, an increase in VAT rates will lead to a reduction in the net prices measured by Istat in the month in which the increase occurs, with the distortion being gradually eliminated in the following months as prices come into line. The consequent erratic behaviour of the official index interferes with the estimation of the seasonal component. Accordingly, the empirical rule described above was used to compute a modified index that excludes changes in indirect taxes; this index differs from the official one only in the short term. However, it is important to note that this treatment does not aim to estimate the "true" incidence of tax changes; whoever (producer or consumer) actually bears the tax burden, the net-of-tax prices are in any case those measured by the index excluding taxes currently published by Istat.

Although changes in indirect taxes are the major source of irregularities, they are not the only one. Other irregularities typically arise from a change in the base year or in prices subject to government control.[9] A more general approach consists in taking account *a priori* of all these changes if they reach a specific magnitude, i.e. a threshold at which the estimation procedure of the seasonal component

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[9] For instance, the reduction of electric charges decided by the Government in July 1996, which contributed about -0.2 points to the monthly rate of change in the CLI, can be regarded as a significant outlier.

would presumably be biased without any external intervention.[10] Since the chief object of the adjustment is to use the data on explicit tax changes and other sources of irregularities in order to treat them as exogenous to the seasonal adjustment procedure, contributions from known and relevant incidental factors should be added back to the seasonally adjusted series.

3. Analysis of the seasonal component: methodological issues

In this section we investigate the type of seasonality faced in the case of the three components of the CLI considered. This issue is empirically relevant since the most popular seasonal adjustment procedures (e.g., the X11-ARIMA method) are designed to remove a particular type of seasonality, a "non-stationary stochastic" seasonality, representable as a non-stationary stochastic process, i.e. a process such that innovations in the seasonal pattern are permanent ("winter becomes summer"). In contrast, when the seasonal shocks are transitory, in other words, their effects vanish in the long run, seasonality is stochastic but "stationary", possibly around a "deterministic" seasonal pattern, which can be simply described by a set of seasonal dummies. In this case, the danger of using such procedures is that they can lead to a loss of information (too much variability is removed from the original series; Maravall, 1996a; Beaulieu and Miron, 1993). In this section, we first test for the presence of non-stationary stochastic seasonality. If there is no evidence of this type of seasonality, there is good reason to test whether seasonality is entirely deterministic or if stochastic seasonality is

[10] In practice, where *ex ante* information is available, such irregularities have been taken into account if their contribution to the monthly rate of change of the CLI is greater than 0.1 points.

present as well. The two cases call for different approaches.

In the following, Δm_t, Δq_t and Δr_t refer to the logarithmic first differences of the corresponding M_t, Q_t and R_t series previously adjusted for outliers using the *ad hoc* procedure just illustrated; first differences have been used since, on the grounds of previous empirical results (Cubadda and Sabbatini, 1995), we assume that the logarithmic levels of aggregate price series have at least one AR root equal to one. Moreover, in order to check for the presence of outliers other than those detected on the basis of *ex ante* information, the TRAMO procedure ("Time Series Regression with ARIMA Noise, Missing Observation and Outliers"; Gomez and Maravall, 1996) has also been used.[11]

We start our empirical investigation by testing for the existence of seasonal unit roots (henceforth, SUR) in the growth rates of the CLI components considered. Let x_t be the first logarithmic difference of a generic CLI component; x_t is said to have a full set of SUR if

$$S(L)\, x_t = \eta_t$$

where $S(L) = \sum_{j=0}^{J-1} L^j$, J is the number of seasons in a year, and η_t follows an ARMA process which does not include any root of the filter $S(L)$ in its MA lag polynomial. If J=4 it is easy to see that

$$S(L) = (1+L)(1+L^2)$$

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[11] Such outliers have actually been identified only for levels of the R series. In particular, two level shift outliers have been detected for the periods 1982.4 and 1983.4.

and hence $S(L)$ has three roots on the unit circle. In particular, the root $L=-1$ generates a non-stationary biannual cycle, whereas the complex conjugate roots $L = \pm i$ are responsible for a non-stationary annual cycle. In the monthly case, that is $J=12$, we have:

$$S(L) = (1+L) \prod_{j=1}^{5} \left\{ 1 - 2\cos(j\pi/6)\, L + L^2 \right\}$$

and hence the root $L=-1$ generates a non-stationary two-month cycle, whereas the pair of complex conjugate roots $L = \exp(\pm ij\pi/6)$, for $j=1,2,\ldots,5$, is responsible for a non-stationary cycle of a period equal to $12/j$.

Note that a series may have unit roots at only some of the seasonal frequencies. Clearly, in this case only a subset of the seasonal cycles will be non-stationary. An important implication of the presence of SUR is that changes in the seasonal pattern are permanent ("winter becomes summer"), whereas if the AR roots associated with seasonal frequencies lie outside the unit circle, the effects of seasonal shocks vanish in the long run.

In order to test for SUR in each of the three series, we used the so-called HEGY test (Hylleberg *et al.*, 1990) for the quarterly series Δq_t and Δr_t, and the test by Franses (1991) for the monthly series, Δm_t. These tests represent extensions of the well-known Dickey-Fuller tests to the SUR case. The version of these tests implemented here compares the existence of each SUR with the presence of a stationary stochastic seasonal cycle plus a deterministic seasonal pattern.

In every cases we found strong evidence against the existence of complex conjugate seasonal unit roots; the null of a non-stationary two-period cycle is not rejected for the Δr_t series, and only marginally rejected for the other series (Table 3).

Considering that there is no evidence of a *full set* of SUR in the three series considered, it is natural to ask if the seasonal fluctuations of these series can be fully captured by deterministic seasonal patterns. In this case, seasonal adjustment consists simply in using dummy variables to remove the seasonal means from the original series. However, if only stationary stochastic seasonality is present in the data, optimal seasonal adjustment involves filtering the original series on the grounds of signal-extraction theory (Pierce 1978; Hillmer and Tiao, 1982). Clearly, these two types of seasonality may coexist in the data, in which case it is necessary both to remove seasonal means and filter.

Some simple tests to discriminate between stochastic and deterministic seasonality were originally proposed by Pierce (1978); here, the modified version suggested by Mills and Mills (1992) is implemented. These tests are based on the assumption that a series x_t can be decomposed into a seasonal component s_t and a non seasonal component n_t:

$$x_t = n_t + s_t .$$

Moreover, these components are both deterministic and stochastic. Hence, we can write:

$$s_t = s_{rt} + s_{dt} \quad , \quad n_t = n_{rt} + n_{dt}$$

where the subscript d indicates the deterministic component and the subscript r indicates the random component. Specifically, we assume that the deterministic components take the forms

$$n_{dt} = \mu_0 + \beta t, \qquad s_{dt} = \sum_{j=1}^{J-1} \mu_j d_{jt}$$

where d_{jt}, for $j=1,2,\ldots,J-1$, are seasonal dummy variables. Finally, the stochastic components follow the processes

$$n_{rt} = \psi_n(L)\varepsilon_\tau, \qquad s_{rt} = \psi_s(L)v_t$$

where ε_t and v_t are white noise processes and $\psi_n(L)$ and $\psi_s(L)$ are possibly divergent polynomials in L. Clearly, the series x_t exhibits a deterministic seasonal pattern if $s_{dt}\neq 0$ and it has stochastic seasonality if $s_{rt}\neq 0$.

Considering that the processes n_{rt} and s_{rt} are not directly observable, we need to place some restrictions in order to identify and estimate these components. Pierce (1978) proposed two assumptions which allow us to solve this identification problem. The first is the "minimal extraction" principle, in other words the variance of the seasonal component must be minimized in order to extract no more than necessary from the original series when removing its seasonality. The second is a "filter preserving" property, that is the seasonal component of the detrended series is obtained by passing the seasonal component of the original series through the same filter used for detrending.

Under these assumptions the test procedure goes as follows. First, the series $\Delta(L)x_t$, where $\Delta(L)$ is a lowfrequency filter such that $\Delta(L)\, n_{rt}$ is stationary,[12] is regressed on a set of seasonal dummies; the resulting residuals u_t^* are then consistent estimates of $u_t = \Delta(L)(s_{rt} + n_{rt})$. Second, these residuals are used to estimate the lag polynomials of the ARMA model

$$\Phi_p(L)(1 - \phi_j L^J)\, u_t^* = \Theta_q(L)(1 - \vartheta_j L^J)\, \varepsilon_t$$

where $\Phi_p(L) = 1 + \Phi_1 L + \ldots + \Phi_p\, L^p$ and $\Theta_q(L) = 1 + \Theta_1 L + \ldots + \Theta_q\, L^q$ are respectively the non-seasonal AR and MA polynomials. Following Pierce (1978) and Mills and Mills (1992), we know that s_{rt}=0 is equivalent to $\phi_J = \vartheta_J = 0$. Hence, ML tests for both the absence of stochastic seasonality and the absence of deterministic seasonality ($\mu_1 = \mu_2 = \ldots = \mu_{J-1} = 0$) can easily be computed.

A practical problem to be solved is the choice of the ARMA model for residuals u_t^*. In order to make this choice, we again relied on TRAMO for automatic identification of the models.[13] Details of the estimated model are reported in Table 4, along with the results of the tests. On the basis of the earlier empirical estimation of the seasonal pattern of the CLI, it is important to note that all the series considered show deterministic seasonality, but not stochastic seasonality.[14]

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[12] According to the results of the zero frequency unit root test reported in Table 3, $\Delta(L) = 1 - L$ was chosen for all the series considered.

[13] Clearly, in order to implement the test, seasonal AR and MA components are included anyway.

[14] For completeness of results, these tests were also performed directly on the aggregate CLI, although it is worth noting that the presence of quarterly components seriously biases the genuine seasonal properties of this series. The results obtained do not differ from those from the other series.

It is important to take into account a potential drawback of the above tests. By checking the significance of the ARMA coefficients at lag J, these tests implicitily ignore the possibility that a series may possess only a subset of the AR and/or MA roots associated with the zero and seasonal frequencies. To clarify this point, consider the case of a quarterly time series generated by the following ARMA(0,2) process:

$$x_t = (1 - \Theta_2 L^2)\varepsilon_t$$

where $\Theta_2 \in (0,1)$. Series x_t clearly has a stochastic stationary annual cycle, which is underscored by a peak in its spectrum at frequency $\pi/2$, although both ϕ_4 and ϑ_4 are equal to zero. This problem can theoretically be solved by running alternative tests in the frequency domain, e.g., by using the statistics suggested by Canova (1996). Unfortunately, non-parametric testing in the frequency domain needs large samples to work properly and hence is not recommended in this context.

However, a possible solution consists in focusing on parametric estimates of the spectra of the series considered. In fact, since we have identified and estimated adequate ARMA models for the stationarized series u_t^* of each variable, we can compute the theoretical power spectra of these models. In particular, the coefficients of the ARMA models fitted are reported in Table 5, together with the usual diagnostic tests, which show no signs of mis-specification for these models. From Table 5, it is clear that both the spectra of the stationarized series of M_t and Q_t are monotonically increasing from frequency zero to π. Thus, these series have a two-month and two-quarter stationary cycle respectively. The form of the spectrum of R_t's stationarized series is plotted in Figure 3.

We can see that there are two peaks: the larger is again located at frequency π, and the smaller is located approximately at frequency 0.473π. Hence, series Δr_t also appears to have a relevant biannual cycle.

Note that the above evidence does not conflict with the results of the Pierce-Mills-Mills tests, which correctly indicate that none of the variables has a full set of stochastic seasonal cycles, whereas the estimated spectra provide evidence of a two-period stochastic cycle in all the series. It is also worth remarking that the evidence derived from the HEGY test of a bi-annual unit root in series Δr_t cannot be considered conclusive. In fact, the estimated ARMA model of Δr_t gives values of the biannual AR and MA roots close to one (0.93 and 0.71 respectively); this is the so-called "near-cancellation" case, in which the HEGY test performs badly due to very serious size distortion (Ghysels, Lee and Noh, 1994). By contrast, in the case of series Δq_t and Δm_t, the results in Table 5 suggest that these cycles are generated only by biannual MA roots and are, therefore, stationary.

To sum up, the results reported here provide clear evidence of a deterministic seasonal pattern in the three components of the CLI. As far as stochastic seasonality is concerned, the only (stochastic) seasonal cycle that appears to exist in the data, once the effects of seasonal dummies have been removed, is a two-period cycle, which is therefore entirely responsible for the high-frequency variability of each CLI component.

The implications of these results for seasonal adjustment of the CLI are very important. The most widely used seasonal adjustment procedure, X11-ARIMA, assumes that the series to be adjusted has a full set of SUR, in other words, non-stationary stochastic seasonality is present (Ghysels and Perron, 1993). This contrasts with the evidence just reported. As already mentioned, in such a situation X11-ARIMA would lead to a loss of information, since too much high-frequency variability would be removed from the original series ("overadjustment"). This is an important point in favour of using methods based on filters derived directly from the characteristics of the time series in question ("model-based" approach), whereas the X11-ARIMA method uses a decomposition procedure that applies a standard moving average filter to all time series, regardless of their nature.

Specifically, we adopt the procedure known as SEATS ("Signal Extraction in ARIMA Time Series"; Gomez and Maravall, 1996), which first identifies an ARIMA model for the series under study and then proceeds to disaggregate it into components (Maravall, 1996a).[15] Since the filter in SEATS is derived directly from the characteristics of the series, it is flexible enough to provide reasonable results when seasonality is fairly stable (although not entirely fixed, otherwise seasonal dummies would be able to capture the seasonal component of the index); this point is extensively discussed in Maravall (1996b). In short, in the extreme case of perfect deterministic seasonality, a seasonal difference is imposed in the ARIMA model so that deterministic seasonality is taken into account, but at the same time the MA seasonal parameter tends to approximate -1 (the actual value is -0.98), so that the seasonal unit roots are nearly cancelled out.

[15] For a detailed comparison of different methods, see Eurostat (1995). It is worth noting that before using SEATS, the TRAMO package should to be used in order to detect outliers.

Although in our case the characteristics of the seasonal pattern of the series considered "naturally" lead us to prefer SEATS, this approach also has more general advantages over the X11-ARIMA. In particular, for the estimated seasonal component it provides estimation standard errors and forecasts (with their associated standard errors), useful tools for short-term monitoring (Maravall, 1996a).[16]

4. Extraction of the seasonal component: empirical results

The three series taken into consideration were adjusted for seasonality using the SEATS procedure. The seasonal adjustment procedure introduced in the first section can be summarized for each series as follows (Table 6). First, outliers are detected and their contribution to the rate of change in the index is removed from the original series.[17] Second, the series obtained is adjusted for seasonality.[18] Finally, the contribution of outliers is added back to the seasonally adjusted series.

Although from a methodological point of view the indirect approach to seasonal adjustment of the CLI introduced in the first section is preferable, in practice it can be helpful to investigate the seasonal pattern of the general index as well (direct approach). The main reason is that the procedure just described calls for disaggregated data, which

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[16] Most of studies that have compared SEATS with X11-ARIMA show that the former method performs better in most real situations (see for example Fiorentini and Planas, 1997; Maravall and Planas, 1997; Planas, 1996 and 1997).

[17] Details of the outliers actually removed on the basis of *ex ante* information are given in Gavosto, Sabbatini and Sestito (1996).

[18] The TRAMO and SEATS packages were adopted using default options, except for the following parameters: LAM=-1; IATIP=1; SEATS=2; INIC=3; IDIF=3 (Gomez and Maravall, 1996).

are not available when provisional figures are released.[19] The CLI seasonally adjusted using the direct approach does not significantly differ from that based on seasonally-adjusted disaggregated data (Figures 4 and 5);[20] hence, it can provide a preliminary assessment of the performance of inflation when provisional figures are released. On the basis of the similiarity between the results obtained following the two different approaches, a more radical solution is to refer directly to the CLI after it has been seasonally adjusted using the direct approach; although this is less "clean" from a methodological point of view, it clearly has very important operational advantages (Table 7 reports the seasonal components estimated on the basis of the direct adjustment).

These results are fairly robust with regard to the choice of the seasonal adjustment method. With reference to the general CLI, Gavosto, Sabbatini and Sestito (1996) and Banca d'Italia (1996) adopted the same type of approach, except that the X11-ARIMA procedure was used and the period considered was 1986-1995. As shown in Figures 6 and 7, differences between the two approaches are not very great. It is also worth noting that the degree of smoothness, a standard empirical criterion for defining a "satisfactory seasonal adjustment", is essentially the same for the two series, and not very different from that in the seasonal adjusted series obtained following Gavosto, Sabbatini and Sestito (1996) and Banca d'Italia (1996, Table 8).[21]

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[19] Moreover, disaggregated data are not available to all users.

[20] The general index has been adjusted for seasonality on the basis of the same set of options used for disaggregated data. For the last two years the general index is slightly different from that published by Istat in that two decimals, instead of one, are considered.

[21] There are no generally accepted (definite) analytical criteria to define a "satisfactory" seasonal adjustment (see for example Sims, 1978; Granger, 1978; Den Butter and Fase, 1991; for confirmation that the different criteria are not really binding, see Van der Hoeven and Hunderpool, 1986). From the point of view of the policy maker, which

A final important point concerns revisions of the seasonal pattern once new observations are added. From the point of view of the policy maker, the fewer revisions required when new data are added, the more desirable a seasonal adjusted series becomes.[22] An analysis was therefore conducted to compare the revision of the seasonal coefficients, estimated directly for the general index, in years: (A) 1992-93 (comparing coefficients estimated respectively over the periods 1981-1993 and 1981-1994); (B) 1993-94 (comparing coefficients estimated respectively over the periods 1981-1994 and 1981-1995). The results reported in Table 9 show how the magnitude of revisions is small in both cases.[23] They also show that, taking a longer time interval than the one considered in Gavosto, Sabbatini and Sestito (1996), the revision of the seasonal component of the general index is on average approximately the same for both X11-ARIMA and SEATS, thus avoiding a possible preference for the former procedure.

5. Conclusion

This paper has focused on the seasonal adjustment of

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is the one adopted in this paper, one often decisive empirical criterion is actually based on the degree of smoothness of the seasonally adjusted series. Intuitively, this requisite is linked directly to the reasons underlying the adjustment, which can be traced to the removal of those effects that (together with incidental factors) prevent the prompt detection of changes in the underlying trend of the series under observation. Two further requisites, equally desirable but partly conflicting with the former, concern the regularity of the seasonal component and the minimization of revisions when new data are added.

22 Although revisions are unpleasant from the point of view of the policy maker, they are "optimal" from a statistical point of view (Maravall, 1997).

23 The ARIMA model selected for each series remains unchanged when one year at time added after 1993, with a slight revision in the estimated parameters.

the cost-of-living index, the most influential measure of Italian inflation. The first part of the paper showed the importance of seasonal variations in explaining the overall variability of the index: in the period 1981-1995, seasonal dummies explain around 20 per cent of the overall variability of the index. It has also been stressed how most of this seasonality is actually spurious, being essentially related to the fact that data on a significant number of the items included in the cost-of-living index are collected quarterly. In order to adjust the cost-of-living index for seasonality it is therefore very important to take account of the different nature of the series involved.

The second part of the paper investigated the characteristics of seasonality in the series involved. It is important to note that different methods must be used, depending on the characteristics of the series considered. The main result of the second section is that the presence of a deterministic seasonal pattern in the components of the CLI is unambiguously detected, whereas the only apparent stochastic seasonal cycle in the data is a two-period cycle, which is thus entirely responsible for the high-frequency variability of each component once seasonal dummies have been removed. This result has very important practical implications for the seasonal adjustment of the series analyzed in the paper. The use in these circumstances of the X11-ARIMA method, which is based on the assumption that the series to be adjusted has a full set of SUR, would induce an "overadjustment", in other words, too much high-frequency variability would be removed from the original series. On the other hand, SEATS is flexible enough to provide reasonable results when seasonality is relatively stable, as the filter is based directly on the characteristics of the series.

SEATS was implemented to adjust the series considered for seasonality. The measure recommended in the paper for monitoring the performance of inflation is the quarterly growth rate of the "indirectly" adjusted series, for two main reasons: first, a significant price change is always considered for all the series; second, it allows us to smooth out irregularities in price changes. A "genuine" monthly signal on inflation can in fact only be provided by the seasonally-adjusted monthly series. In practice, it has been shown that from an empirical point of view the "directly" seasonally-adjusted index provides a close approximation of the seasonal pattern revealed by the "indirectly" seasonally-adjusted measure. Thus, from a practical point of view the former measure can also be used. This is especially useful in the case of the CLI since provisional figures are released for the general index only.

Table 1

WEIGHTS OF THE MAIN COMPONENTS OF THE COST-OF-LIVING INDEX (1)
(percentages)

	base-year				
	1980	1985	1989	1992	1995
Food	32.60	28.41	21.90	20.30	19.68
unprocessed	15.38	10.84	10.05	9.08	8.49
Energy products	6.01	6.18	4.16	3.96	3.73
Non-food non-energy products	33.74	28.87	36.16	39.43	35.66
Services	16.66	21.68	24.77	22.63	25.16
TOTAL goods and services not subject to price controls	89.01	85.14	86.99	86.32	84.23
Goods subject to price controls	1.12	2.43	1.21	1.52	2.61
Public utility charges	5.37	7.84	7.76	8.16	9.78
Rents	4.50	4.59	4.04	4.00	3.38
TOTAL goods and services subject to price controls	10.99	14.86	13.01	13.68	15.77
Quarterly data (including rents) (2)	10.31	14.40	19.29	20.27	18.84
Monthly data (2)	89.69	85.60	80.71	79.73	81.16

Source: Based on Istat data.

(1) Tobacco products and "Medical tests" are not included in the basket; weights are rescaled to 100.

(2) Estimates based on the 1995 frequency of price collection.

Table 2

IMPORTANCE OF SEASONAL DUMMIES IN EXPLAINING THE SHORT-TERM VARIABILITY OF THE COMPONENTS OF THE CLI (1)

R^2 class (share of the variance explained by seasonal dummies)	Ratio of the weight of elementary items included in each R^2 class to the total weight of the aggregate indicated in each column (percentages) (2)						
	Food	Non-food non-energy		Energy	Services (excluding rents)		Total
		monthly (3)	quarterly (4)		monthly (3)	quarterly (4)	
0.10 ≤	8.15	12.88	3.37	-	1.25	-	6.15
0.11 - 0.30	67.40	33.38	15.62	100	51.24	43.19	48.17
0.31 - 0.50	14.02	15.66	7.92	-	37.27	-	16.56
0.51 - 0.70	8.46	14.07	16.59	-	10.24	-	9.91
> 0.70	1.97	24.01	56.50	-	-	56.81	19.21
mean R^2	0.27	0.39	0.56	0.17	0.25	0.42	100
aggregate index R^2 (5)	0.32	0.42	0.89	0.11	0.23	0.32	0.44

R^2 aggregated indices	
including changes in indirect taxes	0.44
monthly	0.32
quarterly (6)	0.73 (0.45)
rents (6)	0.93 (0.78)
excluding changes in indirect taxes	0.48

1) Elementary indices are only available since 1990. Only those included in both the 1989 and 1992 baskets are considered; their total weight (net of tobacco products) is equal to 98.05 and 91.64 per cent respectively. "Medical tests", for which prices are collected every six months, have not been considered (their total weight is equal to 0.13 per cent in the 1992 basket). The weight of each series has been rescaled to total 100 in each base year.
2) For each of the items falling in the aggregates indicated, first the R^2 of a regression of the first difference of each index (log) is computed on 12 seasonal dummies (over the 1990-95 years), and its weight is then added to that of other items falling in the same R^2 class.
3) Series collected monthly.
4) Series collected quarterly, considered at monthly frequencies.
5) R^2 of a regression of each aggregate (taken as a whole rather than item by item) on 12 seasonal dummies.
6) R^2 of a regression of the monthly aggregate relative to the quarterly collected series on 12 seasonal dummies; in brackets is the R^2 of a regression of the quarterly aggregate series on 4 seasonal dummies.

Table 3

TESTING FOR SEASONAL UNIT ROOTS IN THE COST-OF-LIVING COMPONENTS (*)

	Lags	Monthly seasonal frequencies						
		$\pi^{(a)}$	$\pi/2^{(b)}$	$2\pi/3^{(b)}$	$\pi/3^{(b)}$	$5\pi/6^{(b)}$	$\pi/6^{(b)}$	$0^{(a)}$
Δm_t	-	-2.66 (c)	16.79 (c)	13.30 (c)	8.97 (c)	10.02 (c)	9.69 (c)	2.64
		Quarterly seasonal frequencies						
			$\pi^{(a)}$			$\pi/2^{(b)}$		$0^{(a)}$
Δr_t	2, 5		2.47			19.45 (c)		-1.65
Δq_t	-		3.36 (c)			7.62 (c)		-2.69

(*) (i) (a) T-type test ; (b) F-type test for complex conjugate unit roots; (c) rejected at the 5 per cent level. The null is that parameters in the auxiliary regression associated with frequencies indicated are equal to 0, i.e. there is a (seasonal) unit root at the corresponding frequency.

ii) Lags is the number of lagged values of the dependent variable included in the regression.

iii) All series are in the first difference of the log linearized levels, where outliers have been detected and removed using the TRAMO package (Gomez and Maravall, 1996). Note that although level shift outliers have been removed from the level of the series, once the series is differenced these outliers become additive; hence, their removal does not affect the asymptotic test distribution (see Franses and Haldrup, 1994).

iv) The estimation equations include a constant, eleven (three) seasonal dummies, a time trend, and the listed number of lags of the dependent variable.

v) Standard errors are OLS standard errors.

vi) Critical values for monthly series (Franses, 1991): t_{π}=2.65; t_0= 3.24; $F_{\pi/2}$= 5.63; $F_{2\pi/3}$= 5.84; $F_{\pi/3}$= 5.90; $F_{5\pi/6}$ = 5.71; $F_{\pi/6}$= 5.84. Critical values for quarterly series (see Hylleberg *et al.*, 1990): t_{π}=3.08; t_0=3.71; $F_{\pi/2}$= 6.55.

Table 4

TESTING FOR DETERMINISTIC AND STOCHASTIC SEASONALITY IN THE GROWTH RATES OF THE COST-OF-LIVING COMPONENTS (*)

	Non-seasonal ARIMA parameters	ϕ_j	θ_j	F_1	F_2	LB serial correlation test(ii)	JB normality test
Rents	θ_1 = 0.69 θ_2 = -0.54 θ_3 = -0.61 ϕ_1 = 0.92	-0.49	-0.46	3.61 [0.02]	0.00 [0.99]	7.26 [0.70]	0.07 [0.96]
Quarterly data	θ_1 = -0.49	-0.43	-0.26	3.49 [0.02]	0.74 [0.48]	9.27 [0.75]	0.23 [0.89]
Monthly data	θ_1 = -0.77	-0.26	-0.16	11.71 [0.00]	0.71 [0.49]	16.71 [0.21]	0.29 [0.87]

(*) (i) F_1 , F_2 = F-type test for the absence of deterministic and stochastic seasonality, respectively; P-values are given in square brackets.
(ii) Serial correlation test is such that 16 lags are considered for the quarterly series and 24 lags for the monthly series.

ESTIMATED ARMA MODELS
FOR PARAMETRIC SPECTRAL ESTIMATION (*)

	Non-seasonal Arima parameters	LB serial correlation test	JB normality test
Rents	$\theta_1 = 0.71$ $\theta_2 = -0.52$ $\theta_3 = -0.62$ $\phi_1 = 0.93$	6.79 [0.87]	0.12 [0.94]
Quarterly-base collected data	$\theta_1 = -0.48$	16.12 [0.37]	0.06 [0.97]
Monthly-base collected data	$\theta_1 = -0.77$	21.54 [0.55]	1.55 [0.46]

(*) (i) F_1, F_2 = F-type test for the absence of deterministic and stochastic seasonality, respectively; P-values are given in square brackets.
(ii) Serial correlation test is such that 16 lags are considered for the quarterly series and 24 lags for the monthly one.

Table 6

THE MAIN STEPS OF THE CLI SEASONAL ADJUSTMENT PROCEDURE

1. Starting data (1)	M Q (quarterly frequency) R (quarterly frequency)
2. Adjustment for outliers	i) Preliminary adj. on the basis of *ex ante* information (Gavosto, Sabbatini and Sestito, 1996) ii) Automatic outlier detection - TRAMO procedure (Gomez and Maravall, 1996)
3. Seasonal adjustment of M, Q, R	SEATS procedure (Gomez and Maravall, 1996) (2)
4. Seasonal adjusted series	CLI including changes in indirect taxes: preliminary adjustment based on *ex ante* information (step 1) is added back to each seasonally adjusted series
5. Indirectly seasonally adjusted CLI	The quarterly growth rate of the "indirectly" adjusted series is considered in order to monitor underlying inflationary tendencies. A close approximation is provided by the same measure calculated with respect to the "directly" seasonally adjusted index.

1) Disaggregated data for the CLI are only available to us from 1990. In order to consider a longer period of time, quarterly-based collected series and rents have been estimated for the years 1981-1989 on the basis of the corresponding indices for the CPI; the monthly aggregate was obtained as a residual.

2) TRAMO and SEATS packages were adopted using default options, except for the following parameters: LAM=-1; IATIP=1; SEATS=2; INIC=3; IDIF=3 (Gomez and Maravall, 1996).

Table 7

SEASONAL COMPONENT OF THE CLI INCLUDING CHANGES IN INDIRECT TAXES (DIRECT APPROACH) (*)

	Jan.	Feb.	Mar.	Apr.	May	June	July	Ago.	Sep.	Oct.	Nov.	Dec.
1986	0.19	0.15	-0.07	-0.02	-0.05	-0.10	-0.22	-0.21	0.01	0.38	0.10	-0.16
1987	0.18	0.15	-0.06	-0.02	-0.06	-0.10	-0.22	-0.20	0.01	0.36	0.11	-0.15
1988	0.17	0.14	-0.08	0.00	-0.05	-0.09	-0.21	-0.18	-0.01	0.34	0.12	-0.15
1989	0.16	0.14	-0.08	-0.01	-0.03	-0.08	-0.20	-0.18	-0.01	0.33	0.11	-0.15
1990	0.16	0.13	-0.08	-0.01	-0.02	-0.08	-0.20	-0.16	-0.01	0.30	0.11	-0.14
1991	0.15	0.12	-0.07	0.00	-0.02	-0.08	-0.18	-0.15	-0.01	0.28	0.10	-0.14
1992	0.14	0.11	-0.07	0.01	-0.02	-0.06	-0.16	-0.16	-0.02	0.26	0.10	-0.13
1993	0.13	0.10	-0.07	0.00	0.00	-0.06	-0.17	-0.14	-0.01	0.25	0.10	-0.13
1994	0.12	0.10	-0.06	0.00	0.00	-0.05	-0.15	-0.15	-0.02	0.23	0.11	-0.13
1995	0.11	0.10	-0.05	0.00	0.01	-0.06	-0.14	-0.14	-0.02	0.21	0.10	-0.12
1996	0.11	0.09	-0.05	0.00	0.01	-0.06	-0.13	-0.13	-0.02	0.19	0.11	-0.12

(*) Calculated as the difference between the monthly rate of change in the unadjusted and seasonally adjusted series (the seasonal component is here estimated in the period from 1981 to 1996).

Table 8

COMPARISON BETWEEN SEASONALLY ADJUSTED SERIES (1)

	Standard deviation (1990-95)		Mean range of the seasonal component (4) (1986-1995)
	σ_{ra} (2)	σ_{gra} (3)	
Components			
Monthly	2.522	1.796	0.050
Quarterly	1.321	-	0.296
Rents	1.604	-	0.380
General index			
"Indirect" approach	-	1.479	-
"Direct" approach	2.013	1.460	0.053
X11-ARIMA (5)	1.952	1.381	0.118

1) Each series has been adjusted for seasonality with respect to the period 1981-1995.
2) With respect to the rate of change on the previous period, net of the seasonal component, on a annual base.
3) With respect to the three-month rate of change, net of the seasonal component, annualized.
4) Mean, for each month, of ranges of the seasonal components, calculated as the difference between the rate of change on the previous period of the unadjusted series and the rate of change of the seasonally adjusted series.
5) The procedure followed is described in Gavosto, Sabbatini and Sestito (1996), and applied to the period 1981-1995.

Table 9

REVISION OF THE SEASONAL COMPONENT (1)

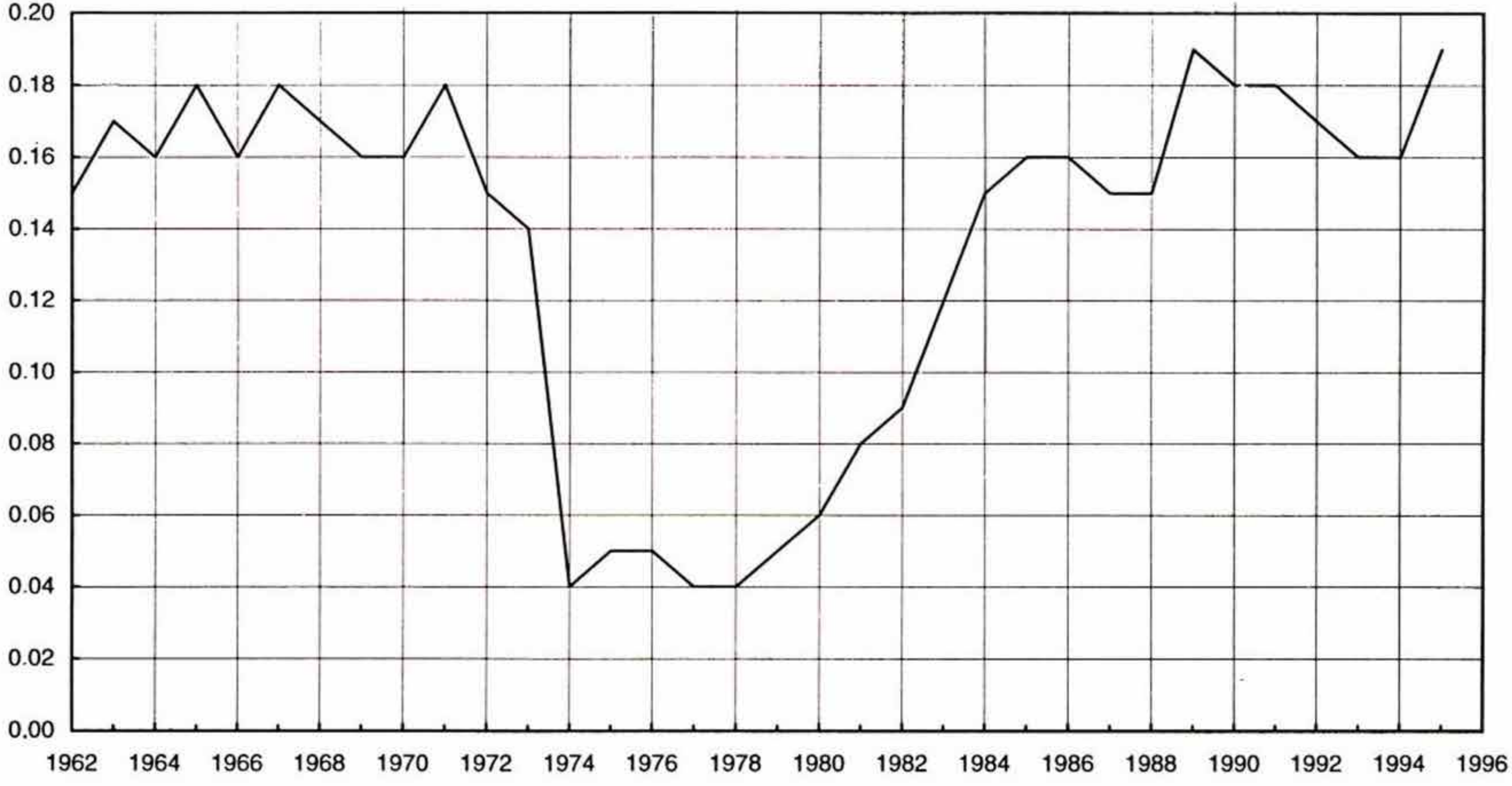
		Monthly series (SEATS)	General index (SEATS)	General index (X11-ARIMA) (2)			Quarterly series (SEATS)	Rents (SEATS)
Jan.	A	0.005	0.012	0.027				
	B	0.012	0.015	0.063				
Feb.	A	0.006	0.008	0.004	1st qtr.	A	0.003	0.120
	B	0.026	0.024	0.073		B	-	0.098
Mar.	A	0.008	0.009	0.011				
	B	0.012	0.002	0.005				
Apr.	A	0.011	0.000	0.004				
	B	0.003	0.002	0.013				
May	A	0.014	0.010	0.026	2nd qtr.	A	0.044	0.033
	B	0.009	0.010	0.016		B	-	0.149
June	A	0.017	0.004	0.022				
	B	0.023	0.013	0.040				
July	A	0.009	0.008	0.019				
	B	0.018	0.015	0.050				
Ago.	A	0.002	0.000	0.007	3rd qtr.	A	0.021	0.023
	B	0.005	0.006	0.032		B	-	0.047
Sept.	A	0.008	0.007	0.006				
	B	0.016	0.010	0.015				
Oct.	A	0.002	0.006	0.036				
	B	0.012	0.012	0.035				
Nov.	A	0.029	0.003	0.059	4th qtr.	A	0.030	0.078
	B	0.004	0.007	0.015		B	-	0.086
Dec.	A	0.023	0.009	0.062				
	B	0.011	0.010	0.007				
Mean	A	0.011	0.006	0.024			0.042	0.063
	B	0.012	0.010	0.030			0.037	0.095

1) A = Mean absolute difference, with respect to the period 1992-93, between the seasonal components estimated respectively on the 1981-1993 and the 1981-1994 periods.; B = Mean absolute difference, with respect to the period 1993-94, between the seasonal components estimated respectively on the 1981-1994 and the 1981-1995 periods.

2) For details of the procedure followed, see Gavosto, Sabbatini and Sestito(1996).

Figure 1

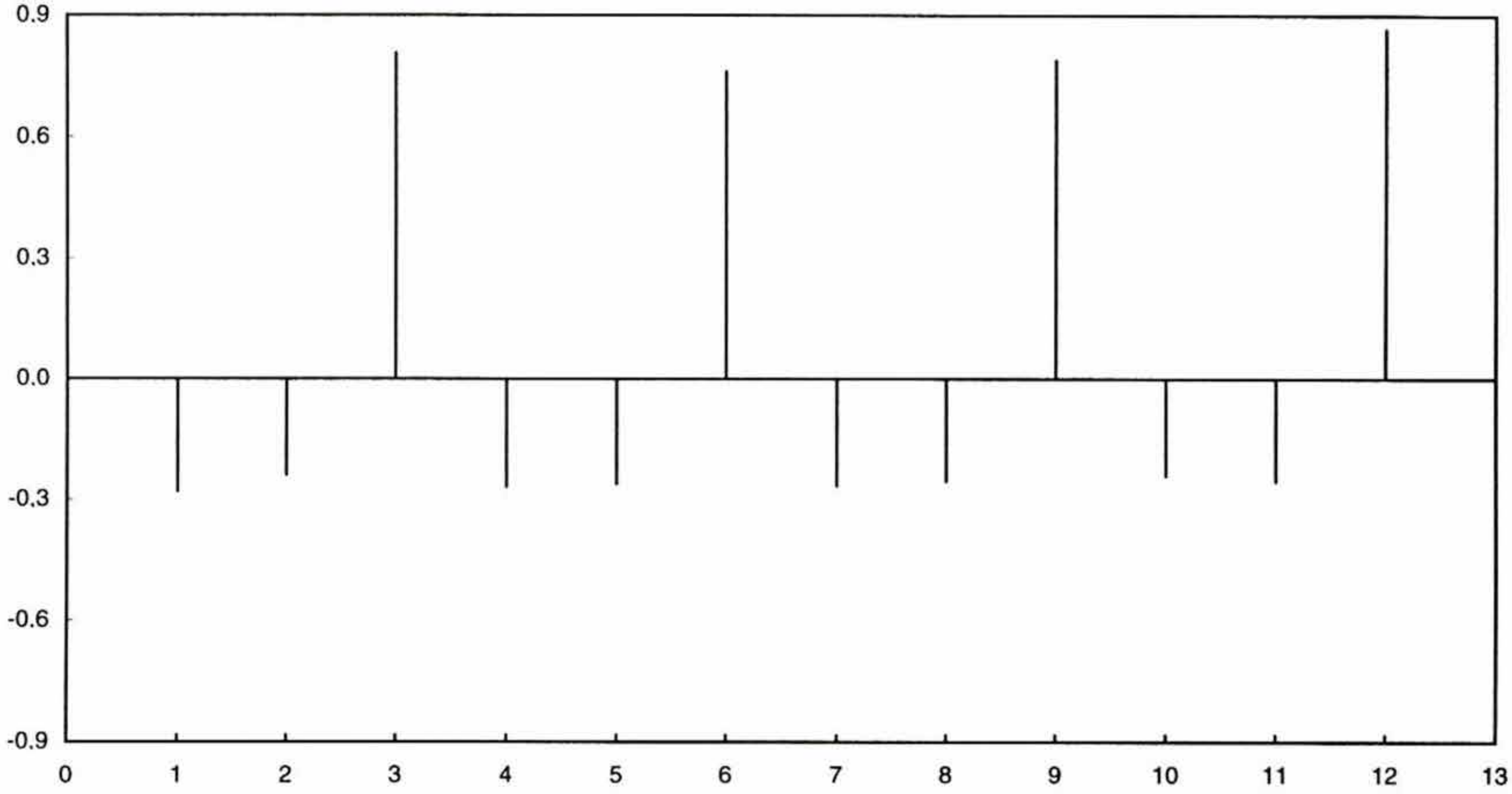
SHORT-TERM CLI VARIABILITY EXPLAINED BY SEASONAL DUMMIES (*)



(*) R^2 of a rolling regression (on 15 years ending in the year indicated) of the first difference of CLI (log) on twelve seasonal dummies.

Figure 2

AUTOCORRELATION FUNCTION OF THE *Q*-SERIES (*)



(*) First difference of a series (log) in which the level of the index is held constant within each quarter.

Figure 3

SPECTRUM OF THE *R* SERIES (STATIONARIZED)

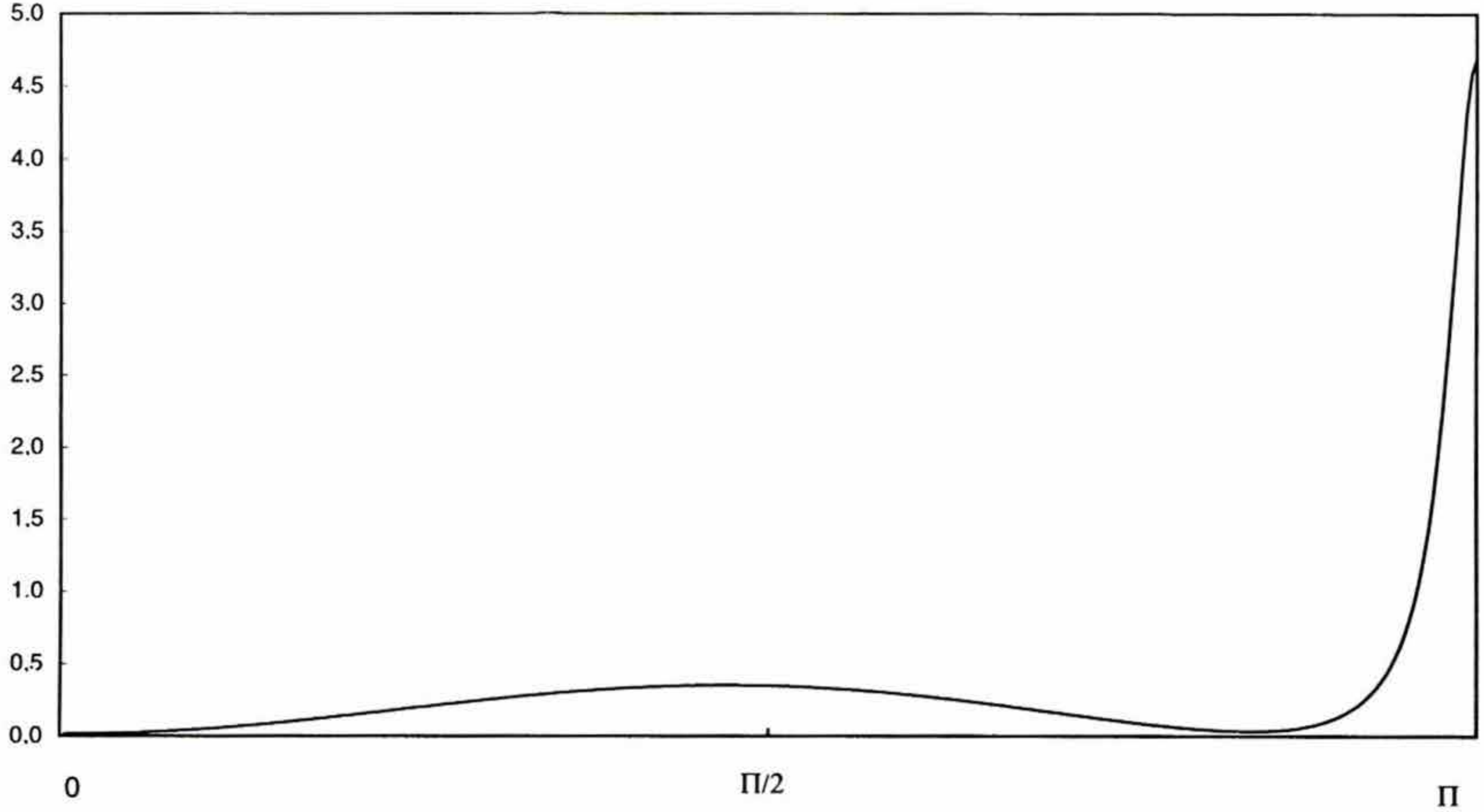
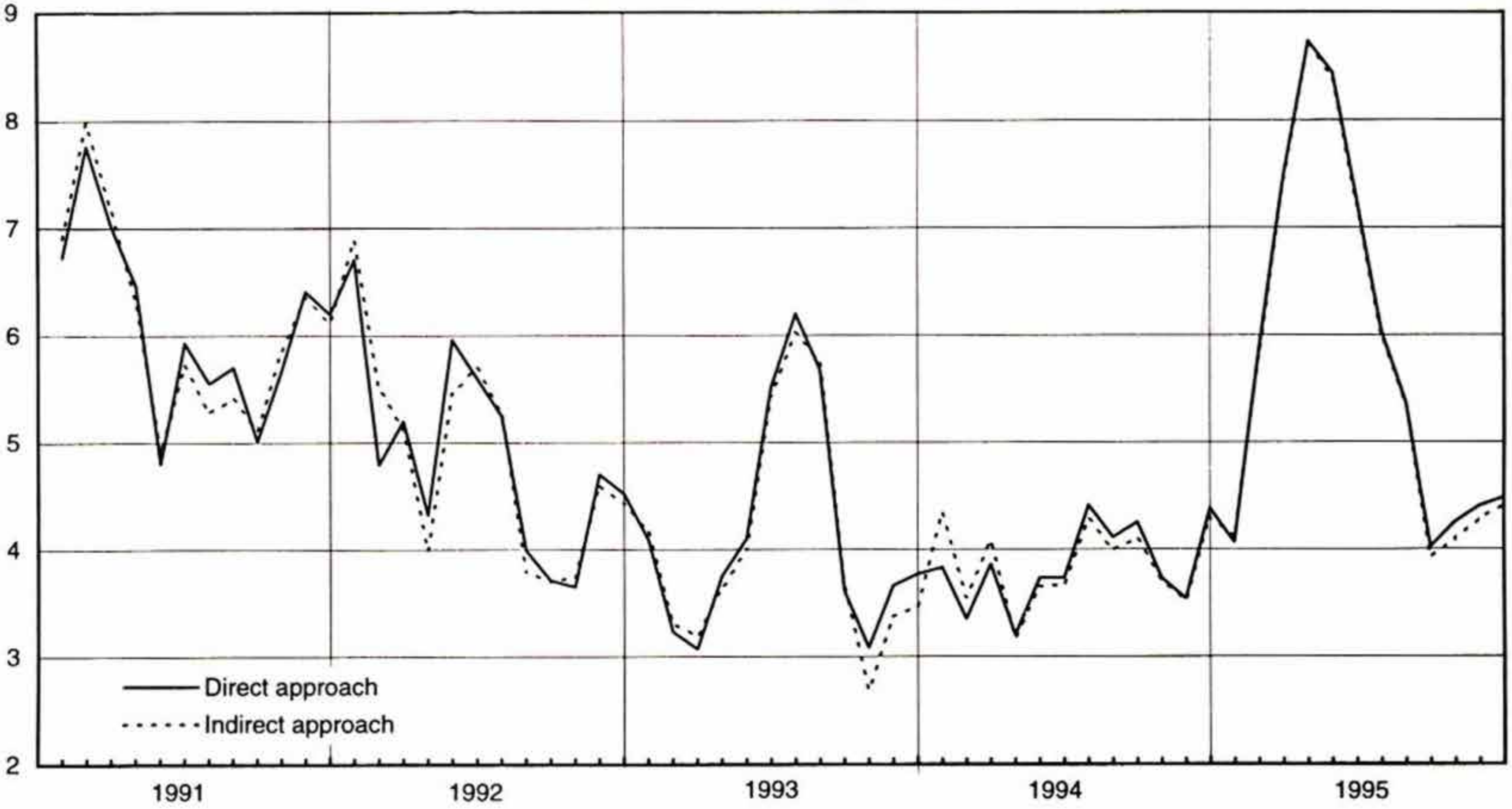


Figure 4

COST-OF-LIVING INDEX SEASONALLY ADJUSTED USING THE SEATS PROCEDURE

(three-month changes, annualized)

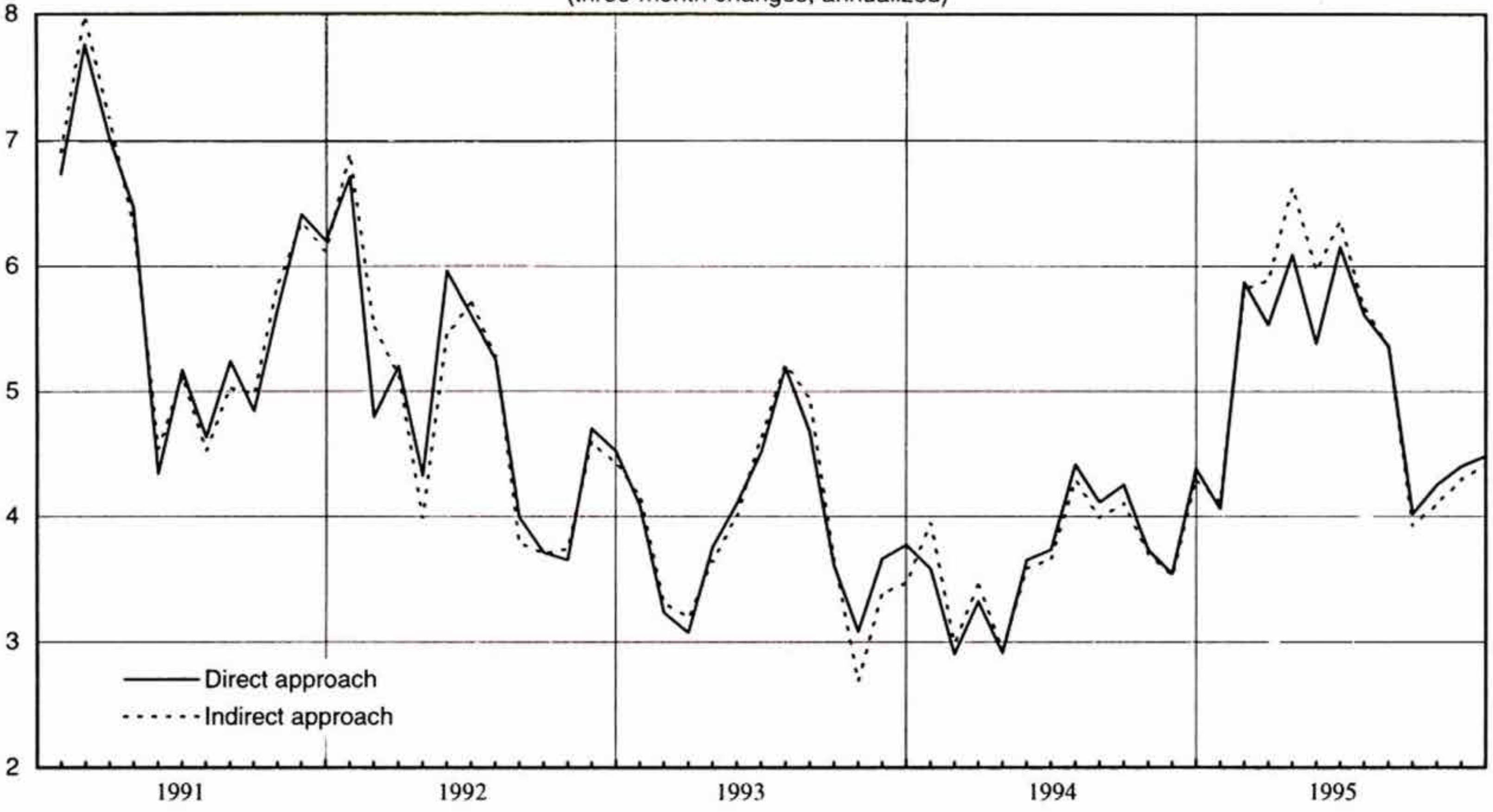


Source: Based on Istat data.

Figure 5

COST-OF-LIVING INDEX EXCLUDING CHANGES IN INDIRECT TAXES, SEASONALLY ADJUSTED USING THE SEATS PROCEDURE

(three-month changes, annualized)

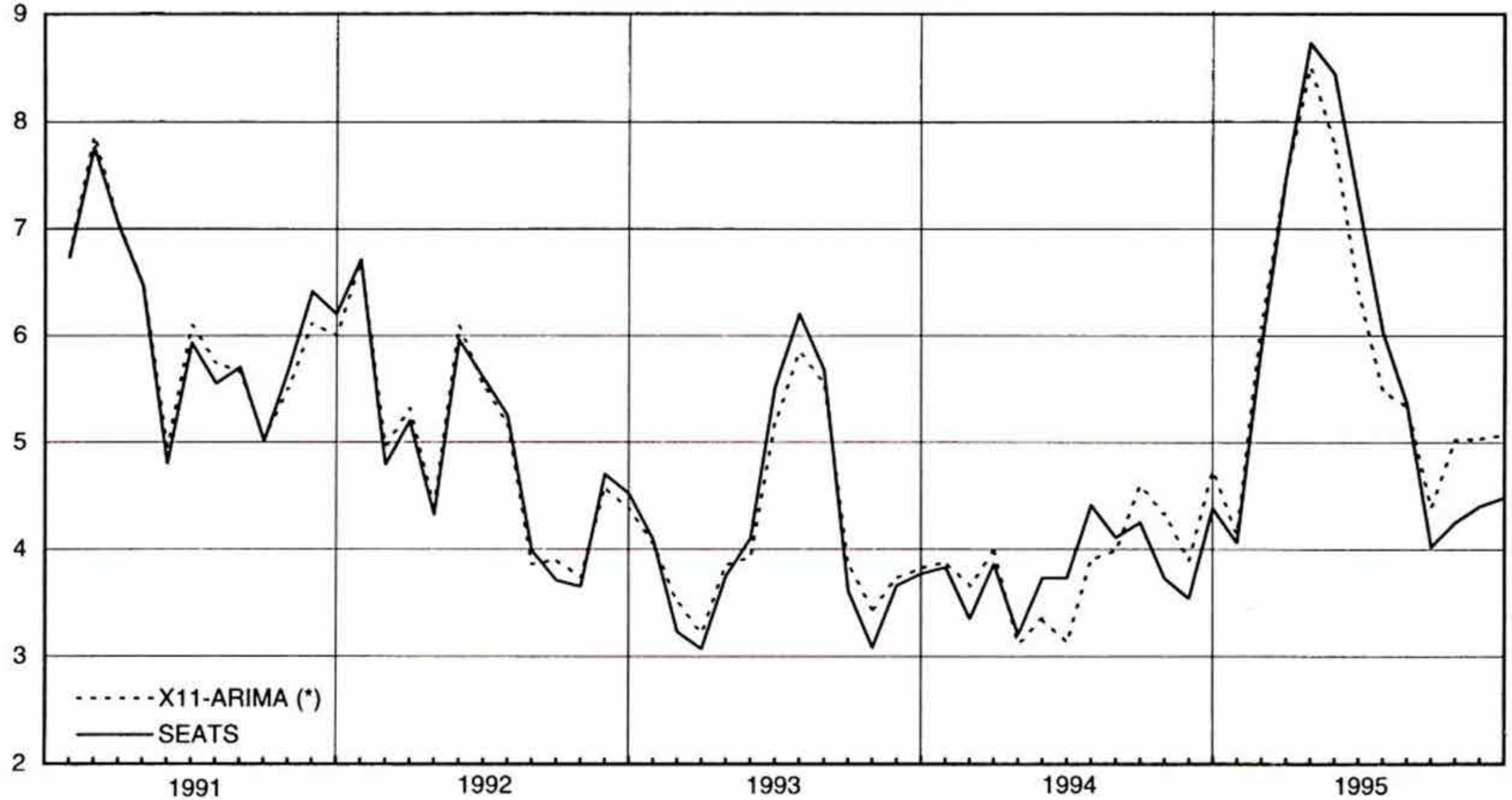


Source: Based on Istat data.

Figure 6

DIFFERENT PROCEDURES FOR SEASONAL ADJUSTMENT OF THE COST-OF-LIVING INDEX - DIRECT APPROACH

(three-month changes, annualized)



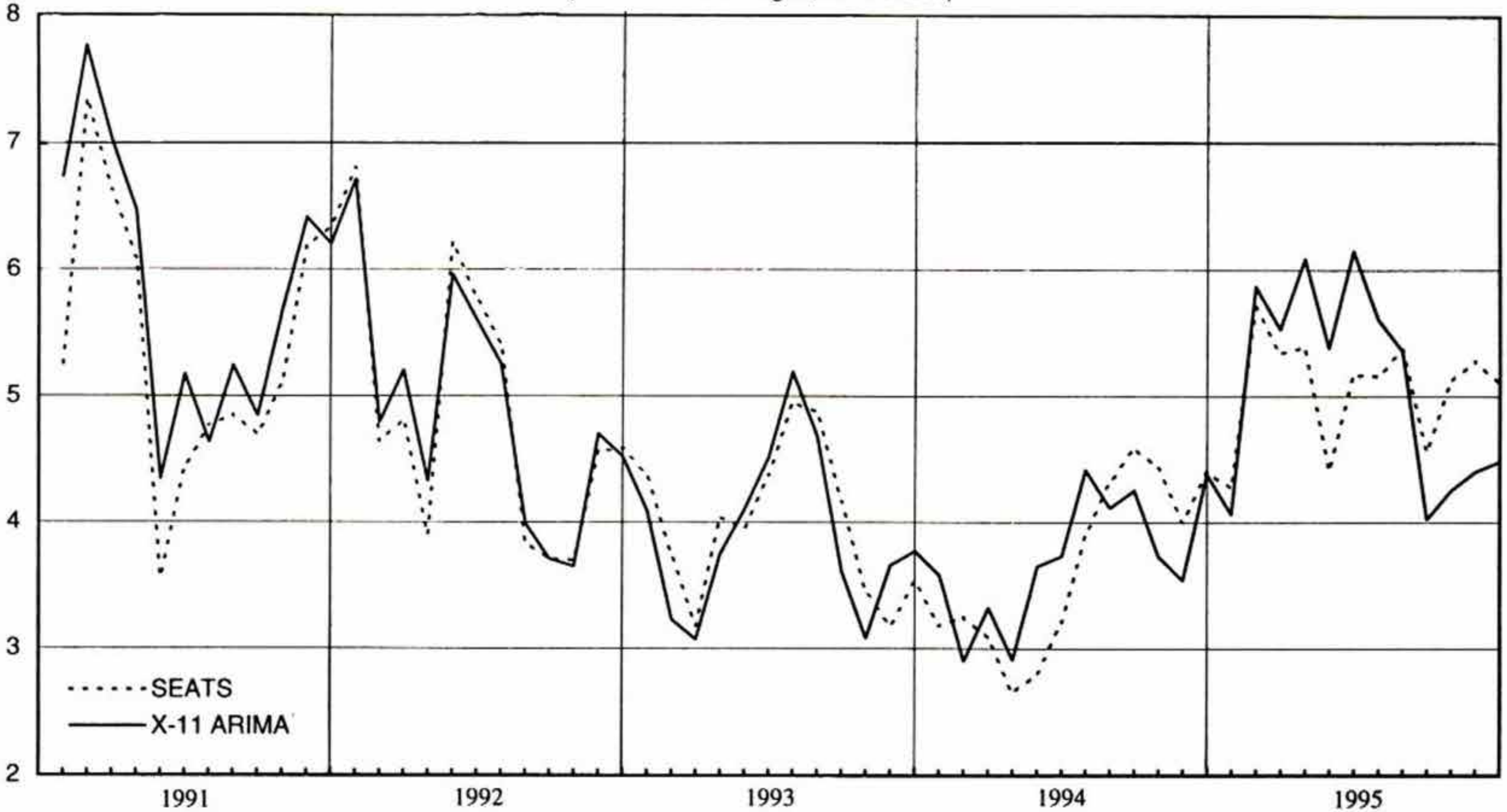
Source: Based on Istat data.

(*) Cost- of-living index seasonally adjusted with the procedure described in Gavosto, Sabbatini and Sestito, 1996.

Figure 7

DIFFERENT PROCEDURES FOR SEASONAL ADJUSTMENT OF THE COST-OF-LIVING INDEX EXCLUDING CHANGES IN INDIRECT TAXES - DIRECT APPROACH (*)

(three-month changes, annualized)



Source: Based on Istat data.

(*) Cost-of-living index seasonally adjusted with the procedure described in Gavosto, Sabbatini and Sestito, 1996.

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