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# *Tax Amnesties and Income Tax Compliance: The Case of Spain*

JULIO LÓPEZ-LABORDA and FERNANDO RODRIGO[*]

***Abstract***

The aim of this paper is to evaluate the long-term impact on Spanish individual income tax (IRPF) compliance of the amnesty measures granted in 1991 within the framework of the 1988–91 income tax reform programme. To that end, we combine time-series techniques with outlier detection methods and the Bai and Perron (1998) test for the endogenous estimation of structural breaks. On the basis of the analysis of the monthly IRPF tax collection series from 1979 to 1998, we find that the amnesty had no effect on tax collection in either the short or the long term. By contrast, we find evidence of the permanent positive impact caused by the legislative and administrative measures linked to the IRPF reform process begun in 1988.

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## I. INTRODUCTION

The phenomenon of tax evasion is present, to a greater or lesser extent, in all the countries of the world. The existence of a high level of tax evasion, which gives rise to numerous distortions in economic efficiency and damages the principle of equity, requires the application of repressive or preventive measures aimed at improving levels of compliance. In this regard, the permanent repression of tax evasion, by way of tax audits and prosecutions in the ordinary Courts of Justice, is occasionally complemented by tax amnesties, which grant taxpayers a second opportunity for a limited period of time to pay the taxes they have evaded in the past. This voluntary regularisation normally implies a partial or total reduction of

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[*]University of Zaragoza.
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the usual penalties that are imposed as a consequence of tax evasion, and even of the interest that has accrued.

During the last few decades, the governments of numerous countries, both developed and less developed, have established tax amnesty measures. Nevertheless, the typology of these programmes has been extremely varied: we have seen the implementation of both temporary and permanent amnesties; on some occasions, tax evaders who have already been detected by way of inspection have been allowed to participate in such amnesties, whilst on others this possibility has been expressly prohibited; some countries have offered amnesties for very specific types of fraud, with others legislating for the totality of the tax regimes currently in force and effect.

Obviously, the granting of a tax amnesty is not an action that is simple to defend, given the interaction of numerous factors that exercise an unknown influence over its possible effects. Its apparent advantages, such as the relatively rapid recovery of tax liabilities and the inclusion of new taxpayers in the tax authorities' records, have to be weighed against the harmful effects that such measures might have on normally honest individuals.

Despite the frequent use that governments have made of tax amnesties, the literature on this topic is still in a relatively early stage of development. Thus, theoretical works that analyse the economic impact of regularisation programmes have been presented as an extension of the conventional modelling of the individual decision to evade income tax, introduced by Allingham and Sandmo (1972) and which is based on expected utility maximisation schemes. Participation in a tax amnesty can be analysed under this type of approach if the view is taken that the individual who regularises his position is only seeking to correct his initial decision to evade. Individuals regularise their tax situations because, quite apart from the phenomenon of tax amnesty, there might also have been simultaneous changes in the tax environment. These can take the form of: expectations of greater enforcement efforts (Alm and Beck, 1990 and 1991; Stella, 1991; Macho-Stadler, Olivella and Pérez, 1993); an immunity provision against tax inspection in exchange for a fixed payment during the amnesty period (Cassone and Marchese, 1995; Franzoni, 1996); or a change in the level of individual income (Andreoni, 1991; Graetz and Wilde, 1993; Marceau and Mongrain, 2000).

This conventional approach has been revised by others, such as Malik and Schwab (1991) in the adaptive utility framework or Alm and Beck (1990) in prospect theory, with these latter approaches offering alternative explanations for individual behaviour in situations of uncertainty and for the specific patterns of conduct observed in amnesties granted in a variety of countries.

As regards the empirical aspect, the results obtained following the granting of various tax amnesties have been evaluated either on the basis of information supplied by the databases and official statistics of different tax authorities, or by way of econometric analyses that have specified structural models or applied time-series methods. On other occasions, the scarcity of empirical data has made laboratory experiments a more advisable approach, where the participants are invited to make tax decisions in the presence of an amnesty.

This empirical research has essentially focused on three aspects. First, a series of works have highlighted the role that tax amnesties can play in favouring the transition to tax environments characterised by a more rigorous prosecution of fraud (Mikesell, 1986; Parle and Hirlinger, 1986; Uchitelle, 1989; Alm, McKee and Beck, 1990; Alm and Beck, 1991). Secondly, other authors (Fisher, Goddeeris and Young, 1989; Das-Gupta and Mookherjee, 1995) have emphasised the limited contribution of tax amnesties in the incorporation of new taxpayers. Thirdly, a group of papers have been devoted to considering the impact of amnesties on long-term compliance (Alm, McKee and Beck, 1990; Pommerehne and Zweifel, 1991; Alm and Beck, 1993; Das-Gupta and Mookherjee, 1995).

With respect to this third group, the literature reveals a significant heterogeneity as regards both the size and the estimated sign of these long-term effects. This heterogeneity can be explained in part by the difficulty of comparing the experiences being analysed. However, there would appear to be no doubt that time-series techniques, as demonstrated by Alm and Beck (1993) for the 1985 Colorado amnesty, represent a useful and appropriate tool when seeking to identify the permanent consequences of amnesties for compliance.

The present paper falls within the line of research followed by this third group. Specifically, its aim is to evaluate the long-term impact on the tax compliance of the Spanish individual income tax (*Impuesto sobre la Renta de las Personas Físicas*, hereafter referred to as IRPF) of the tax amnesty granted in 1991 as part of the 1988–91 tax reform measures. To that end, we have chosen to use time-series techniques, estimating both univariate (ARIMA modelling — Box and Jenkins (1976)) and multivariate (intervention analysis — Box and Tiao (1975)) models.

In our econometric analysis, the variable that is the subject of study is the monthly IRPF real tax collection series corresponding to the period 1979–98. The study is limited to the IRPF, given that the 1991 amnesty was a response to a problem of compliance that was particularly associated with this tax. We use the tax collection data as a proxy for IRPF tax compliance: any positive (negative) net effect of the amnesty on compliance should translate into an increase (decrease) in tax collections.

In our view, the paper offers a number of interesting contributions. For example, and to the best of our knowledge, it is the first piece of research that has been carried out on this topic in Spain and can be added to the still limited number of works in the literature that have their origins in other countries. Furthermore, we introduce a novel econometric instrument in the analysis — namely, the Bai and Perron (1998) test — whose application, combined with methods for detecting outliers, reinforces the interpretation of the results obtained by way of time-series techniques. Following the joint use of these procedures and instruments, the main conclusion that emerges is that the 1991 amnesty had no consequences on subsequent compliance in either the short or the long term.

We are conscious that the principal virtue of the technique employed — that is to say, the limited amount of data that is required — is also its principal defect, in that it omits from consideration a variety of other factors that likely affect tax collections. Therefore, an obvious extension to this work would be to test its results against those coming from a structural model that, in line with that developed by Das-Gupta and Mookherjee (1995), explains IRPF tax collection on the basis of a set of variables that includes the granting of the amnesty and the changes derived from the reform process applied to that tax. These methodological options are, in our view, equally valid and potentially complementary.

The rest of the paper is organised as follows. In Section II, we offer a brief description of the significant reforms made to the IRPF between 1988 and 1991 and the tax regularisation programme implemented in that latter year. Section III is dedicated to the effects of the 1991 tax amnesty on long-term compliance, with consideration being given to the techniques employed and the process followed in the econometric exercise, as well as to the results obtained from it. Section IV closes the paper with a review of the main conclusions.

## II. REFORM OF THE IRPF AND THE 1991 TAX AMNESTY

Following the death of General Franco in 1975, Spain entered into a process of transition towards a democratic regime of government. As part of this process, and in the context of a thoroughgoing reform of the tax system, 1979 saw the introduction of a modern individual income tax, with a synthetic character, an extensive base, a progressive tax schedule and personal and incentive-based tax credits.

During the more than twenty years in which it has remained in force and effect, the IRPF has undergone numerous changes. Amongst these, probably the most important was the reform programme implemented during the period 1988–91, as a result of a ruling of the Spanish Constitutional Court handed down in 1989, which declared the obligatory joint taxation of families to be unconstitutional.

The reform was implemented with the following time sequence. The tax schedule was modified in 1988, especially for the highest incomes: the marginal tax rates for taxable incomes of between approximately 8 and 13 million pesetas were reduced, whilst the marginal rates for taxable incomes higher than this latter amount were increased.

The IRPF Provisional Reform Law was introduced in 1989 in order to bring the law into line with the judgement of the Constitutional Court. The main element of this reform was that the individual replaced the family as the taxpaying unit. By contrast, no substantial changes were made in the determination of the tax base. This reform also applied to incomes obtained during 1988.

Finally, in 1991, a law was introduced to consolidate in one statute the changes that had been introduced since 1988.[1]

Additionally, during the period 1989–90, a set of administrative measures were implemented aimed at improving the tax control of specific forms of capital incomes through the extension of the tax-withholding mechanism and the selective application of tax inspections.

In 1991, the tax reform was complemented with a regularisation process — a true tax amnesty — whose aim was the disclosure of undeclared incomes held in the form of very specific assets.[2] In the Spain of the 1980s, part of undeclared income was channelled through assets that were not subject to tax withholding at source and were of a nature that did not require registration. The majority of these assets had a private character, although, somewhat paradoxically, there was one of a clearly public nature — namely, Spanish Treasury Notes (*Pagarés del Tesoro*). The maintenance of a situation whereby the government itself issued securities carrying a low rate of return, in exchange for guaranteeing their complete fiscal opacity, represented a contradiction with a tax system that was purportedly inspired by principles of equity and generality.

With the aim of bringing this situation to an end, two regularisation mechanisms were implemented, with the regularisation period running from June to December 1991. The first of these was aimed at those taxpayers — individuals or corporations — who held funds in the above-mentioned Spanish Treasury Notes. In response, the government issued another asset, described as Special Public Debt, which carried a nominal return of 2 per cent and which could be exchanged for the Treasury Notes. However, the end of opacity could not be brought about in a brusque manner, given that the holders of the Treasury Notes had invested their money under the expectation, created by the government itself, that the interest received would be net of tax. As a result, the subscription for the new asset was totally anonymous, in such a way that the opacity of its holders was maintained. The identity of these holders could only become known to the government on the ordinary redemption date of the Special Public Debt — that is to say, in June 1997 — and by this time the evaders could benefit from the prescription period and thus avoid liability.

The second pardoning mechanism took the form of a provision that already existed in Spanish law — namely, the voluntary regularisation of any evaded

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[1]Legal stability was to be one of the characteristic features of the IRPF until 1998. The only relevant reform implemented during this period was the reduction in the taxation of capital gains introduced in 1996. For a detailed description of the recent evolution of the IRPF and of the Spanish tax system, see Lasheras and Menéndez (1998).

[2]A very generous tax amnesty had already been granted in 1977 to favour the application of the tax reform implemented under the newly democratic regime.

taxation through the filing of a supplementary tax return. Here, the novel aspect was that such declarations filed during the statutory period — June to December 1991 — did not carry any late-payment interest or penalty whatsoever.

Together with these two mechanisms, a body of legislative reforms and specific administrative measures were put into effect, with the aims of increasing the attraction of participating in the amnesty and of making the future cost of evasion more onerous. Amongst these measures, attention should particularly be drawn to the following: first, the organisational improvements introduced to increase the efficiency of the Spanish Revenue Service; secondly, the commitments — publicised through an aggressive advertising campaign — to intensify the tax inspections of those who chose not to participate in the regularisation, together with the indication that the amnesty represented a final opportunity for the taxpayer to regularise his position with the tax authorities; and, thirdly, the strengthening of the penalty regime. In summary, the amnesty was accompanied by those measures to which the theoretical and applied literature attributes positive effects over the behaviour of tax evaders (Alm and Beck, 1991; Graetz and Wilde, 1993).

The regularisation processes that we have just described obviously present peculiar features on the basis of comparative experience. However, in our view, they possess the requisites that characterise an authentic tax amnesty, in that they suppose the non-application of the penalties provided for in the case of non-compliance with tax obligations under the terms and during the time period specified in the regularisation legislation.

Participation in this amnesty was officially described as massive. However, if we compare these data with the relevant tax magnitudes, the modest nature of the response becomes clear. Tax collection from supplementary tax returns amounted to just 0.8 per cent of the ordinary tax collection corresponding to the 1990 tax year (compared with 0.86 per cent in 1989 and 0.71 per cent in 1991). In the particular case of the IRPF, a total of 36,257 supplementary tax returns were presented, which regularised an average amount of 548,915 pesetas, whilst the approximately 11 million taxpayers who complied with their obligations declared average amounts of 302,920 pesetas in the 1990 tax year and 324,816 pesetas in the 1991 tax year. When seeking to evaluate these figures, account should be taken of the fact that the regularisation could cover various earlier tax years.

Furthermore, the regularisation of the Treasury Notes reached levels corresponding to 63 per cent of these securities.

However, rather than offering various approximations of the immediate results of these regularisation procedures, our interest is in determining whether the 1991 amnesty had any permanent effect, whether positive or negative, on IRPF compliance, and this is the question we consider in the following section.

## III. EFFECTS OF THE 1991 TAX AMNESTY ON LONG-TERM COMPLIANCE

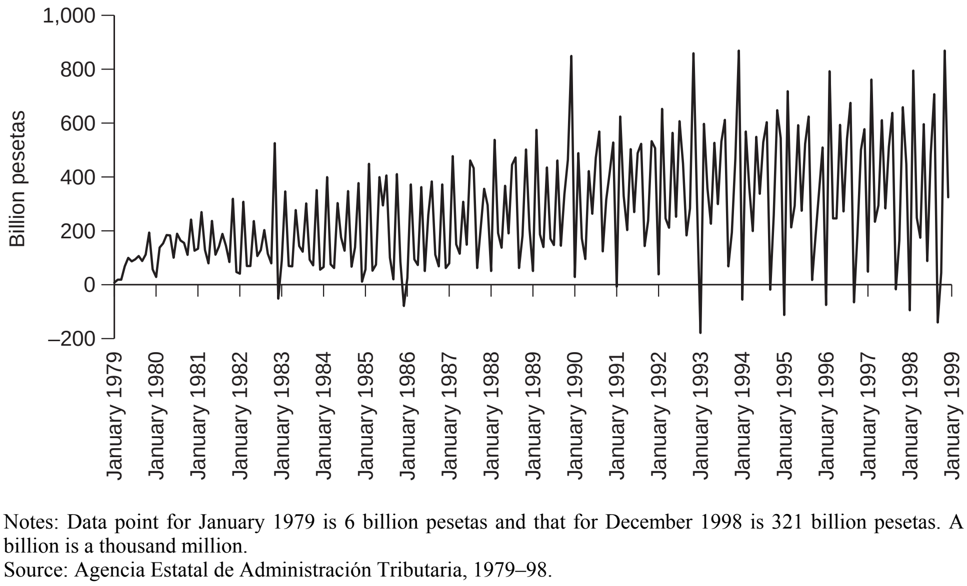
### *1. Description of the Techniques Employed*

Although a variable such as tax collection can, in principle, be affected by a large group of factors (whether or not controlled by the tax authorities), our work sets out to identify and quantify the variations in that variable caused by regularisation — that is to say, by what we have chosen to describe as the 'amnesty effect'. A priori, an amnesty could have both positive and negative effects on the tax collection figure. That is to say, it could influence taxable income in two different ways: positively, if it brings more income and taxpayers into the tax net, or negatively, if it leads to a deterioration in compliance on the part of individuals who normally meet their tax payments.

The choice of time-series techniques as a procedure to identify the long-term effects of the 1991 amnesty can be justified for various reasons. First, such techniques allow for hypothesis tests on the possible impact of a specific measure on the level or trend of a series, which, in our case, is the monthly real tax collection series for IRPF (1979–98), as reflected in Figure 1. An additional advantage of this approach is that it allows us to work with a very limited set of information, with the said series being the only data required for the application of such methods. This series includes a sufficient number of observations, in the

FIGURE 1

**Monthly IRPF Real Tax Collection Series**




Notes: Data point for January 1979 is 6 billion pesetas and that for December 1998 is 321 billion pesetas. A billion is a thousand million.
Source: Agencia Estatal de Administración Tributaria, 1979–98.

periods both prior and subsequent to the amnesty, for the Box–Jenkins methodology to be reliably applied and, following its application, to observe possible differentiated behaviour within the tax collection figures.

Our first concern has been to detect and estimate the ‘amnesty effect’ itself. It can be expected that if this effect is permanent, then the amnesty will cause a break in the series, with tax collection showing a different pattern of behaviour before and after this event. This allows us to formulate our null hypothesis $H_0$ — namely, that the amnesty breaks the series.

In order to test for this, we have used the Box–Jenkins (1976) methodology for estimating ARIMA models. It is precisely this estimation that gives us information on the possible variation in the explanatory model of the series following the amnesty. If such variation is observed, this constitutes a first indication that the amnesty caused a structural break in the series. Two other instruments that are available to test for this are the application of the Chow (1960) test of structural permanence, and a forecast-quality exercise designed for that purpose. This set of instruments allows us to identify two different stochastic structures, corresponding to the periods before and after the granting of the amnesty.

Following this, we have carried out an intervention analysis, which permits the affirmation that the amnesty was not the event that gave rise to the break in the time series. This result, together with the detection of outliers and the Bai and Perron (1998) test, allows us to formulate the alternative hypothesis $H_a$ — namely, that the IRPF reform breaks the series. The reiteration of the Box–Jenkins methodology and of the intervention analysis results in us being able to confirm this alternative hypothesis.

In what follows, we will briefly consider each of the instruments used and the results obtained.

## *2. Identification and Estimation of the ‘Amnesty Effect’ Using Parametric Time-Series Methods*

### (a) Application of the Box–Jenkins Methodology to the Tax Collection Series

Parametric time-series methods, or the univariate Box–Jenkins methodology, start from the past movements of a variable in order to forecast its future behaviour. The typical structure of these models takes the form

$$(1) \qquad y_t = f(y_{t-1}, y_{t-2}, ...) + u_t ,$$

where $y_t$ is the series whose behaviour is being analysed and $u_t$ is a white noise variable.

This methodology is an iterative procedure whose aim is to determine the ARIMA($p$,$d$,$q$) × ARIMA($P$,$D$,$Q$)$_s$ model that is susceptible to have generated the series in question and which, therefore, represents the behaviour of that series. This procedure requires that the chosen model passes each of the four stages proposed by the methodology — namely, identification, estimation, diagnostic checking and forecasting of the time series.[3]

We have selected three periods for the analysis: from January 1979 to December 1998, which covers the whole period during which the IRPF was in force and effect; from January 1979 to May 1991, the period prior to the granting of the amnesty; and from June 1991 to December 1998, the period subsequent to the granting of the amnesty. The stochastic processes selected for each of these periods are as follows:

- complete period (1/79–12/98) — ARIMA$(0,1,1)_{12}$, with the inclusion of an independent term:

  (2) $\left(1-L^{12}\right)y_t = \mu + \left(1-\Theta_1 L^{12}\right)u_t$ ;
- period prior to the amnesty (1/79–5/91) — ARIMA$(1,1,0)_{12}$, with the inclusion of an independent term:

  (3) $\left(1-\Phi_1 L^{12}\right)\left(1-L^{12}\right)y_t = \mu + u_t$ ;
- period subsequent to the amnesty (6/91–12/98) — ARIMA$(1,0,0)(0,1,1)_{12}$:

  (4) $\left(1-\phi_1 L\right)\left(1-L^{12}\right)y_t = \left(1-\Theta_1 L^{12}\right)u_t$ .

These specifications confirm the existence of stochastic structures that are different for each of the periods being analysed. Thus, the first test procedures for structural permanence in the tax collection series would appear to indicate the presence of a time structure that is different for the periods prior and subsequent to the amnesty.

We now apply a structural change test, specifically the Chow (1960) test, taking the time of the granting of the amnesty as the point of reference. In our case, the value of $F$ is 10.39, larger than the critical point corresponding to $F_{\varepsilon=0.05}(K,T–2K) = 3$. Thus, the null hypothesis of serial stability is rejected, and this instrument would also appear to confirm the existence of different series in the periods prior and subsequent to the granting of the amnesty.

The last stage of the Box–Jenkins methodology, centred on forecasting the future values of the series on the basis of the estimated ARIMA structure, can also have an interpretation orientated towards our objective — namely, that if with the model estimated for the pre-amnesty period, we are not able to predict reasonably the real tax collection that is subsequently observed, then we can say that there remains evidence of a structural break in the series.

However, before turning to this forecasting exercise, it is appropriate to carry out a detection analysis of the possible outliers which, although present in the

[3]For a detailed explanation of this methodology, see Franses (1998).

series in question, are impossible to know a priori.[4] The detection of outliers and the proper treatment of them contribute towards improving time-series modelling, in that these unknown external events are capable of altering the structure of the statistics habitually employed in the identification stage. Even when such an identification procedure is adequate, these events can also alter the estimations of the parameters. Furthermore, as Chen, Liu and Hudak (1990) indicate, if the presence of outliers is adjusted for, this improves the quality of the forecasts obtained.

Furthermore, and as we will consider later and in more detail, the detection and specification of these outliers in our exercise can contribute towards explaining specific patterns of behaviour detected in the tax collection series, and thus help in the overall understanding of its evolution.

The different types of outliers analysed in the literature are essentially the following: first, the additive outlier (AO), which is an external event or effect that affects the series at just one moment in time; secondly, the innovative outlier (IO), which is an event whose effect spreads in accordance with the ARIMA model of the process, in such a way that it affects each of the observations of the series subsequent to the appearance of that event; and, finally, the level shift (LS), which is an event that has permanent effects over the series once it has occurred (Aznar and Trívez, 1993).

The outliers found in the different periods, together with their quantification and corresponding t-value, are set out in Table 1. The detection of these outliers was made by way of the iterative procedure proposed by Chang and Tiao (1983).

If we estimate once again the ARIMA processes selected earlier for each of the periods, but now introducing the outliers we have identified, the explanatory capacity of these models improves. This is reflected in the new estimations of the standard deviation of the residuals, $\hat{\sigma}_u$, shown in Table 1. That is to say, the models estimated following this procedure improve significantly in informative content, in such a way that more precise forecasts can be expected.

Thus, once the outliers detected in the period prior to the amnesty are included in the specification of that period — that is to say, in expression (3) — the resulting model takes the form

$$(5) \qquad y_t = \mu + \sum_{i=1}^{5} w_{it} I_t^{t_0} + \sum_{i=1}^{2} w_{it} I_t^{t_0} \frac{1}{\left(1-\Phi_1 L^{12}\right)\left(1-L^{12}\right)} + \frac{u_t}{\left(1-\Phi_1 L^{12}\right)\left(1-L^{12}\right)}.$$

[4]Put another way, time series are frequently influenced by external events that can have an impact on their normal behaviour. The existence of these events is revealed when observing the data outside the ordinary behaviour of the variable although, in the first instance, we do not know what has caused them to happen and, as a result, the period in which they began.

TABLE 1

**Detection of Outliers and Their Quantification in the Estimation of the Different Series Periods**

**ARIMA(0,1,1)$_{12}$ (and independent term) (period: 1/79–12/98)**

| *Observation* | *Value* | *t-value* | *Type*[a] |
|---|---|---|---|
| August 1985 | –255,628.04 | –3.78 | AO |
| October 1985 | 279,956.18 | 4.02 | AO |
| November 1985 | –300,496.09 | –4.04 | AO |
| July 1989 | –297,668.80 | –4.12 | AO |
| December 1989 | 630,592.08 | 6.49 | IO |
| November 1992 | 342,533.90 | 4.43 | AO |
| December 1993 | 400,031.21 | 4.96 | AO |
| $\hat{\sigma}_u =$ 105,730 (without outliers) | | $\hat{\sigma}_u =$ 76,814 (with outliers) | |

**ARIMA(1,1,0)$_{12}$ (and independent term) (period: 1/79–5/91)**

| *Observation* | *Value* | *t-value* | *Type*[a] |
|---|---|---|---|
| November 1982 | 208,813.38 | 3.90 | AO |
| August 1985 | –263,971.85 | –4.30 | AO |
| October 1985 | 277,879.59 | 3.96 | AO |
| November 1985 | –274,165.77 | –4.23 | AO |
| December 1987 | 272,548.92 | 3.97 | IO |
| July 1989 | –322,410.25 | –3.95 | AO |
| December 1989 | 555,180.26 | 5.84 | IO |
| $\hat{\sigma}_u =$ 106,970 (without outliers) | | $\hat{\sigma}_u =$ 64,380 (with outliers) | |

**ARIMA(1,0,0)(0,1,1)$_{12}$ (period: 6/91–12/98)**

| *Observation* | *Value* | *t-value* | *Type*[a] |
|---|---|---|---|
| December 1993 | 345,464.00 | 4.33 | AO |
| $\hat{\sigma}_u =$ 95,637 (without outliers) | | $\hat{\sigma}_u =$ 89,046 (with outliers) | |

[a]Assuming that $z_t$ is the series free of outliers and that $y_t$ is the observed series, an additive outlier (AO) is an event that affects the series at just one moment in time, $t_0$: $y_t = z_t + wI_t^{t_0}$, where $I_t^{t_0}$, which is equal to 1 if $t = t_0$ and to 0 if $t \neq t_0$, is a pulse variable that represents the presence or absence of the outlier in the period $t_0$, and $w$ is the size of the immediate and sole effect of that outlier.

An innovative outlier (IO) reflects an event whose effect spreads in accordance with the identified ARIMA model and that affects all the observed values after its starting point. Its representation is as follows: $y_t = z_t + \frac{\theta(L)}{\phi(L)} wI_t^{t_0}$, where $I_t^{t_0}$ is the same pulse variable as defined earlier, and where $\frac{\theta(L)}{\phi(L)}$ reflects the ARMA($p,q$) of the series. If $\frac{\theta(L)}{\phi(L)} = \psi(L) = 1 + \psi_1 L + \psi_2 L^2 + \ldots$, we can conclude that the effect of the IO that happens in $t = t_0$ over $t_0+j$, for $j > 0$, is equal to $w\psi_j$, where $w$ is the initial effect and $\psi_j$ is the $j$th coefficient of the polynomial $\psi(L)$.

Using (5), we can forecast the post-amnesty period. To that end, we calculate both the one-period-ahead prediction and the error associated with this prediction. This prediction error is defined as $e_T(1) = y_{T+1} - y_T(1)$, where $y_T(1)$ is the prediction calculated in period $T$ for the following period and $y_{T+1}$ is the true value of the series in period $T$+1.

In order to obtain the one-period-ahead predicted tax collection series and the corresponding prediction error, in our exercise we start from the estimation of (5) with informative base up to $T$ = 149 (May 1991) and forecast tax collection for the period $T$+1 = 150. Repeating this exercise with $T$ = 150,...,239, we obtain the series we are looking for. The prediction error series shows that this error increases with the passage of time, and therefore the quality of the forecasts worsens.

We can use some non-parametric prediction-quality indicators to confirm with greater precision this progressive deterioration in the degree of accuracy of the predictions. One possible estimation of this prediction quality is given by the calculation of the mean square prediction error, *MSPE*, which takes the following form:

$$(6) \qquad MSPE = \frac{1}{H}\sum_{i=0}^{H-1} e_{T+i}^2(1), \quad T = 149, \ldots ,$$

where $e_{T+i}(1)$ is the prediction error as it has been defined and $H$ is an arbitrary number of observations which act as a base in order to calculate this average (in this exercise, $H$ = 12). If we calculate this indicator at different times in the predicted period, we can note how its final possible estimations are significantly superior to those produced at the beginning of the period: specifically, the *MSPE* of $T$ = (228,...,239) exceeds that of $T$ = (149,...,160) by some 201 per cent.[5]

The prediction exercise therefore provides new evidence pointing towards structural non-permanence, based on the inadequacy of the model estimated in the pre-amnesty period to forecast what occurred in the post-amnesty period.

[5]An additional indicator of this progressive deterioration in the capacity of the pre-amnesty model to explain the subsequent period is given by the evolution of the value of the standard deviation of the disturbance in the model. The estimation of the model with an informative base up to May 1991 presents a standard deviation of the residuals of $\hat{\sigma}_u$ = 64,098, whilst with an informative base up to November 1998, this increases to $\hat{\sigma}_u$ = 86,728. The value of $\hat{\sigma}_u^2$ plays a very important role in obtaining the forecast's confidence intervals, in that it is equal to the estimated variance of the one-period-ahead prediction error. With a pre-fixed level of significance $\varepsilon$, the confidence interval of $y_{T+1}$ is equal to $y_T(1) \pm N_{\varepsilon/2}\hat{\sigma}_u$. Therefore, the informative content of the predictions made on the basis of the pre-amnesty model declines with the passage of time, with the region of plausible values progressively increasing.

(b) Intervention Analysis

If we review the results obtained up to this point, we can appreciate that, whilst we have identified two different stochastic structures, corresponding to the periods prior and subsequent to the granting of the amnesty, we still do not have econometrically conclusive proof that is sufficient to attribute the structural break to the exclusive presence of the regularisation. In this sense, and as a complementary procedure, intervention analysis represents a new approach for studying the impact of an amnesty.

When we were considering the analysis of the outliers, we indicated that their characterisation was given by the a priori lack of knowledge that the analyst has of both the cause and the time of appearance of these external events. By contrast, intervention analysis allows us to evaluate the effect of events that also have an influence on the time series and whose time of appearance is now known.

In this sense, the 'amnesty' intervention could be characterised as an additive effect on tax collection, with a possible beginning in June 1991 and which, a priori, could have either a positive or a negative sign, a more or less sharp beginning and a temporary or a permanent duration. Our lack of knowledge of the profile of this effect requires that it be introduced into the ARIMA model estimated for the complete tax collection period — expression (2) — by reference to different typologies.

Intervention models can be classified by the starting point of the event (sharp or gradual) and by its duration (permanent or temporary). The modelling is basically carried out by way of dummy variables, whose interaction with other parameters gives rise to the different profiles of the intervention's impact (Aznar and Trívez, 1993).

In the first column of Table 2, we present the different modelling used to test for the 'amnesty effect' in this intervention analysis, as well as the stochastic structures that were finally specified.

The impact of the amnesty will be reflected in any of these models by the coefficient $w_1$ (corresponding to the dummy associated with the intervention). For models I to IV, this impact has been tested for with a commencement date of $t_0$ = June 1991, July 1991, ..., December 1991 — that is to say, each one of the months of the regularisation period established by law — in this way reflecting the possibility that the impact of the amnesty might have begun to show itself with a certain delay.

The table contains the results corresponding to two further models: first, model V, which tries to reflect whether the amnesty had a significant isolated effect in any of the months during which it was in force and effect; and, secondly, model VI, which, given the form of the step variable introduced, which takes the value 1 in the months June–December 1991 and 0 for the rest, reflects whether the impact was limited to the six months of the regularisation period.

TABLE 2

**Results of the Estimation of the Amnesty Effect in the Intervention Models**

| *Modelling the amnesty effect*[a] | *Starting point of the amnesty effect* | *Estimate of the amnesty effect (t-value in parentheses)* |
|---|---|---|
| Model I<br>Sharp starting point and permanent duration<br>$\left(1-L^{12}\right)y_t = \mu + w_1\left(1-L^{12}\right)S_t^{t_0} + \left(1-\Theta_1 L^{12}\right)u_t$ | June 1991 | $w_1$ = 7,112 (0.24) |
| | July 1991 | $w_1$ = 4,716 (0.16) |
| | August 1991 | $w_1$ = 2,711 (0.09) |
| | September 1991 | $w_1$ = 2,177 (0.07) |
| | October 1991 | $w_1$ = –1,413 (–0.05) |
| | November 1991 | $w_1$ = 1,697 (0.06) |
| | December 1991 | $w_1$ = 5,224 (0.18) |
| Model II<br>Gradual starting point and permanent duration<br>$\left(1-\delta L\right)\left(1-L^{12}\right)y_t$<br>$= \mu + w_1\left(1-L^{12}\right)S_t^{t_0} + \left(1-\delta L\right)\left(1-\Theta_1 L^{12}\right)u_t$ | June 1991 | $w_1$ = 1,942 (0.26)<br>$\delta$ = 0.76 (3.55) |
| | July 1991 | $w_1$ = 1,490 (0.20)<br>$\delta$ = 0.76 (3.55) |
| | August 1991 | $w_1$ = 1,226 (0.16)<br>$\delta$ = 0.77 (3.43) |
| | September 1991 | $w_1$ = 926 (0.12)<br>$\delta$ = 0.77 (3.53) |
| | October 1991 | $w_1$ = 616 (0.08)<br>$\delta$ = 0.77 (3.52) |
| | November 1991 | $w_1$ = 560 (0.07)<br>$\delta$ = 0.77 (3.50) |
| | December 1991 | $w_1$ = 205 (0.03)<br>$\delta$ = 0.77 (3.49) |
| Model III<br>Sharp starting point and temporary duration<br>$\left(1-\delta L\right)\left(1-L^{12}\right)y_t$<br>$= \mu + w_1\left(1-L^{12}\right)I_t^{t_0} + \left(1-\delta L\right)\left(1-\Theta_1 L^{12}\right)u_t$ | June 1991 | $w_1$ = 26,530 (0.46)<br>$\delta$ = 0.78 (3.85) |
| | July 1991 | $w_1$ = 22,053 (0.38)<br>$\delta$ = 0.78 (3.83) |
| | August 1991 | $w_1$ = 17,768 (0.31)<br>$\delta$ = 0.78 (3.80) |
| | September 1991 | $w_1$ = 20,821 (0.36)<br>$\delta$ = 0.78 (0.20) |
| | October 1991 | $w_1$ = 1,171 (0.02)<br>$\delta$ = 0.80 (3.42) |
| | November 1991 | $w_1$ = 30,830 (0.58)<br>$\delta$ = 0.76 (4.42) |
| | December 1991 | $w_1$ = 50,972 (0.88)<br>$\delta$ = 0.80 (4.68) |

TABLE 2 continued

| | | |
|---|---|---|
| Model IV<br>Gradual starting point and temporary duration<br>$\left(1-\delta_1 L-\delta_2 L^2\right)\left(1-L^{12}\right)y_t$<br>$=\mu+w_1\left(1-L^{12}\right)I_t^{t_0}+\left(1-\delta_1 L-\delta_2 L^2\right)\left(1-\Theta_1 L^{12}\right)u_t$ | June 1991 | $w_1$ = 27,897 (0.35)<br>$\delta_1$ = 0.51 (0.77)<br>$\delta_2$ = 0.10 (0.17) |
| | July 1991 | $w_1$ = 24,038 (0.31)<br>$\delta_1$ = 0.31 (0.43)<br>$\delta_2$ = 0.28 (0.44) |
| | August 1991 | $w_1$ = 11,139 (0.14)<br>$\delta_1$ = 0.32 (0.44)<br>$\delta_2$ = 0.27 (0.41) |
| | September 1991 | $w_1$ = 25,414 (0.32)<br>$\delta_1$ = 0.47 (0.72)<br>$\delta_2$ = 0.14 (0.23) |
| | October 1991 | $w_1$ = –10,925 (–0.14)<br>$\delta_1$ = 0.46 (0.67)<br>$\delta_2$ = 0.12 (0.19) |
| | November 1991 | $w_1$ = 1,883 (0.02)<br>$\delta_1$ = 0.36 (0.50)<br>$\delta_2$ = 0.22 (0.35) |
| | December 1991 | $w_1$ = 44,363 (0.58)<br>$\delta_1$ = 0.54 (0.76)<br>$\delta_2$ = 0.13 (0.20) |
| Model V<br>Intervention of the effect limited to only one month of the regularisation period<br>$\left(1-L^{12}\right)y_t=\mu+w_1\left(1-L^{12}\right)I_t^{150-156}+\left(1-\Theta_1 L^{12}\right)u_t$ | June 1991 | $w_1$ = 23,495 (0.26) |
| | July 1991 | $w_1$ = 20,514 (0.22) |
| | August 1991 | $w_1$ = 4,780 (0.05) |
| | September 1991 | $w_1$ = 35,898 (0.39) |
| | October 1991 | $w_1$ = –30,647 (–0.33) |
| | November 1991 | $w_1$ = –38,184 (–0.42) |
| | December 1991 | $w_1$ = –1,216 (–0.01) |
| Model VI<br>Intervention of the effect limited to the totality of the regularisation period<br>$\left(1-L^{12}\right)y_t=\mu+w_1\left(1-L^{12}\right)S_t^{150-156}+\left(1-\Theta_1 L^{12}\right)u_t$ | June 1991 | $w_1$ = 2,148 (0.06) |

[a]If our aim is to reflect a permanent effect of the amnesty, then the dummy variable used in the modelling takes the form of a step variable defined as $S_t^{t_0}$ = {0 prior to the event and 1 after it}, with $t_0$ being the starting point of the intervention. If, by contrast, we wish to reflect a temporary impact, then we use a pulse variable defined as $I_t^{t_0}$ = {1 in the period in which the event happens and 0 in the remaining periods}, with $t_0$ being the starting point of the intervention.

Table 2 also contains the estimates, under the different models, of coefficient $w_1$ and of the corresponding t-values. As the main conclusion, we can note that none of the models has been able to capture a significant amnesty effect (t-value of $w_1$ always lower than 2). Indeed, if we consider that by way of model V we can capture isolated impacts in each of the months during which the amnesty was in force and effect, we cannot even maintain that this amnesty had a contemporaneous impact on taxpayers.

### *3. Joint Reading of the Results. Detection of Alternative Permanent Events: Analysis of Outliers and Application of the Bai and Perron Test*

In the light of these results which, a priori, have contradictory implications, we are obliged to draw an interpretation of them that is consistent. We can first note that one part of these results leads us to affirm that there are clear indications of a structural break in the IRPF tax collection series. However, it is also the case that the absence of structural stability in the series must be translated into the presence of a dummy variable whose impact on the series is permanent from the time of its appearance and which, as a consequence, can be identified as the cause of the break. The results obtained from the intervention models indicate that this variable is not present in any of the months of the regularisation period — that is to say, the 'amnesty' event does not break the series.

In reality, the intervention models indicate that the amnesty did not even have an impact on tax collection during the months affected by the regularisation period. Thus, a first conclusion of this exercise is that we cannot demonstrate that the 1991 regularisation had any impact, either positive or negative, on subsequent IRPF compliance.

In an attempt to complete the exercise, and given our finding that the tax collection series does not behave in the same way throughout the period under study, we should now try to identify what other events might have significantly modified the behaviour of the tax collection figures. To that end, we will consider both the results obtained from the outliers identification procedure and the additional evidence that might be provided by the Bai and Perron (1998) test.

The results obtained from the outliers identification procedure can be particularly useful. In this regard, and from a study of the contents of Table 1, which presents the outliers detected in the complete estimation period, we can note the presence of an IO-type outlier which, given its estimated sign and the time of its appearance, had a positive and permanent impact on tax collection from December 1989. By contrast, the other outliers, by virtue of being of the AO type, only had an impact on tax collection in the respective months in which they appeared.

Similarly, the procedure proposed by Bai and Perron (1998) can supply additional evidence in this identification process. These authors developed a methodology that allows us to test for the presence of structural breaks in a series and which indicates the period in which they appear, further demonstrating its consistency when estimating the number of breaks and their time of appearance.

This test has the following form. First, taking all the available observations of the series, $y_t$, we test for the existence of a structural change using a *sup-F* test. This is based on the differences between the following sums of squared residuals:

(7) non-restricted regression: $y_t = \mu + \delta DU_t + u_t$ ,

(8) restricted regression: $y_t = \mu + u_t$ ,

where $DU_t$ is equal to 1 when $t$ exceeds the time of the break and 0 otherwise. The $F$ statistic is calculated repeatedly, testing with successive times of break, and, calculated in this way, the $F$ statistic with a highest value becomes the *sup-F*. On the basis of the values tabulated by Bai and Perron, we can determine, with a given level of significance, whether or not this value of the *sup-F* indicates the presence of a structural break. Having calculated a date of significant break, the entire series is divided into the two resultant subperiods and the procedure applied to each of them. If new significant breaks appear, the subdivision process continues until none of the resultant subperiods contains additional break points.

The application of this methodology to the complete IRPF tax collection series has produced the results set out in Table 3. As we can observe, four different periods have been detected endogenously, with each distinct period being associated with a change in the mean of the series. If we focus on the evolution of these mean values, we can see how the move to a different period is always associated with the achievement of higher tax collections.

The last of these breaks, which took place in June 1990, could have a direct relationship with the IO-type outlier detected in December 1989. Furthermore, this change in tax collection that has been identified by both methods might well

TABLE 3

**Results of the Application of the Bai and Perron (1998) Test to the IRPF Tax Collection Series**

| *Tax collection periods detected* | *Estimated average for the period (million pesetas) (t-value in parentheses)* |
|---|---|
| 1st tax collection stage: January 1979 to September1982 | 127,075<br>(4.39) |
| 2nd tax collection stage: October 1982 to December 1986 | 177,732<br>(6.47) |
| 3rd tax collection stage: January 1987 to May 1990 | 284,520<br>(9.29) |
| 4th tax collection stage: June 1990 to December 1998 | 370,911<br>(19.04) |

be linked to a greater extent with the IRPF reform begun in 1988 than with the 'amnesty effect' itself, with this being the case for the following reasons.

First, and as we said earlier, the IRPF reform substituted the earlier family-based tax system for one that was based strictly on the individual. This change could have had quite a significant effect on the decision of married women to earn income. In this regard, the literature has found a clearly negative relationship between the IRPF marginal tax rates and the rates of female participation in the labour market.[6]

Recently, Badenes (2000), following the methodology of Feldstein (1995), has estimated the behaviour effects of the 1988–91 IRPF reform, quantifying the response elasticities of the taxable income of second-income receivers in the face of changes in the percentage of income net of taxes. The significant magnitude demonstrated by this elasticity could provide support for the connection between the reform and the change in level observed in tax collection between the end of 1989 and mid-1990.[7] In this regard, account should be taken of the fact that the new system of individual taxation could only induce taxpayers to obtain a higher income from 1989 onwards — that is to say, from the year in which this system was introduced. Whilst it is true that individual taxation also applied to the 1988 tax year, taxpayers had already made their decisions with respect to income under the terms of the earlier family-based taxation legislation.

In addition to this incentive effect, other factors could also have contributed to this increase in tax collection — for example, the intensification of tax inspection activity on the part of the authorities during 1989–90, or the increase in actual income upon which this tax was levied, in that the 1989 and subsequent tax years coincided with a period of strong economic growth in Spain. However, it is more difficult to attribute some influence to the change in the tax schedule in that, as we have already indicated, this benefited some taxpayers but prejudiced others.

In conclusion, both the outlier detection exercise and the application of the Bai and Perron test point to the changes undergone by the IRPF in 1988 as a possible event that gave rise to the break in the series. By contrast, the other break points selected by that test (October 1982 and January 1987) do not find similar empirical support in the detection of outliers.[8]

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[6]For example, García, González-Páramo and Zabalza (1989) have estimated the elasticity of female participation with respect to wage variations at a value of 1.6, and the elasticity of the number of hours offered at 1.9. For Blanco (1992), these values were 2 and 2.8 respectively. This last author has also calculated both elasticities with respect to men, with the values being 0.4 and 0.9 respectively.

[7]Specifically, the value offered by Badenes for this elasticity was in the range 1.17 to 2.09 for 1989, with the estimates being progressively higher for subsequent years.

[8]Nevertheless, mention should be made of some factors that might lie behind these other two jumps in tax collection. As regards the increase associated with October 1982, this could be explained: first, by the extension of taxable income resulting from the increase in income; secondly, by the role played by the fiscal drag effect; and, finally, by the successive increases in nominal tax rates during the period 1982–84. As regards the permanent tax collection gains of 1987, these might have resulted, first, from the positive repercussion on IRPF of the expansive phase of the economic cycle that had begun in Spain in 1986 and, secondly, from the

These results allow us to formulate an alternative hypothesis of a break in the monthly IRPF collection series — namely, that the level of collection can be explained by the reform to the IRPF begun in 1988. In what follows, we will therefore test this new hypothesis, repeating the procedure followed in subsection III.2.

*4. Testing the Alternative Hypothesis: The IRPF Reform as the Explanatory Event of the Serial Break*

Taking the tax collection impact detected as an IO-type outlier (December 1989) as the break point of the series, and using the time-series model estimation procedure, our first objective is to demonstrate the presence of two distinct structures — that is to say, before and after this event.

In this case, the structures specified and estimated for each of these two new subperiods are

- prior period (1/79–11/89) — ARIMA(3,0,0)(0,1,1)$_{12}$, with the inclusion of an independent term and $\phi_1 = \phi_2 = 0$:

  (9) $\left(1-\phi_3 L^3\right)\left(1-L^{12}\right) y_t = \mu + \left(1-\Theta_1 L^{12}\right) u_t$;
- subsequent period (12/89–12/98) — ARIMA(0,1,1)$_{12}$:

  (10) $\left(1-L^{12}\right) y_t = \left(1-\Theta_1 L^{12}\right) u_t$.

Therefore, the estimation of the different models is a first indication of the break as from December 1989.

The Chow test, applied to the new tax collection subsamples, also provides evidence of this lack of series stability. Effectively, the value of the *F* statistic in this case is $F$ = 23.78 > $F_{\varepsilon=0.05}(K,T\text{–}2K)$ = 3, so that the null hypothesis of structural permanence is rejected.

In the final analysis, the bad predictive performance of the subsequent series, on the basis of the modelling of the prior period, can also be viewed as further evidence of this lack of structural permanence. Thus, and as we have justified in subsection III.2(a)), we incorporate the new outliers detected in the first period (Table 4) into expression (9) and obtain the following specification:

$$(11) \qquad y_t = \mu + \sum_{i=1}^{4} w_{it} I_t^{t_0} + \sum_{i=1}^{3} w_{it} I_t^{t_0} \frac{\left(1-\Theta_1 L^{12}\right)}{\left(1-\phi_3 L^3\right)\left(1-L^{12}\right)} + \frac{\left(1-\Theta_1 L^{12}\right) u_t}{\left(1-\phi_3 L^3\right)\left(1-L^{12}\right)}.$$

intensification of anti-fraud campaigns on the part of the tax authorities that had begun in 1985 and which were based on selected tax inspections of specific individuals and companies.

TABLE 4

**Detection of Outliers and Their Quantification in the Estimations of the New Series Periods**

**ARIMA(3,0,0)(0,1,1)$_{12}$ (with independent term and $\phi_1 = \phi_2 = 0$ ) (period: 1/79–11/89)**

| *Observation* | *Value* | *t-value* | *Type*[a] |
|---|---|---|---|
| November 1982 | 211,250.94 | 4.15 | AO |
| July 1985 | 243,332.39 | 3.94 | IO |
| August 1985 | –249,438.16 | –3.74 | AO |
| October 1985 | 246,985.80 | 3.91 | AO |
| November 1985 | –274,905.23 | –3.48 | IO |
| December 1987 | 250,350.54 | 3.81 | IO |
| July 1989 | –281,335.79 | –3.43 | AO |
| $\hat{\sigma}_u =$ 87,012 (without outliers) | | $\hat{\sigma}_u =$ 57,288 (with outliers) | |

**ARIMA(0,1,1)$_{12}$ (period: 12/89–12/98)**

| *Observation* | *Value* | *t-value* | *Type*[a] |
|---|---|---|---|
| November 1992 | 322,601.55 | 4.15 | AO |
| December 1992 | –275,234.57 | –3.33 | IO |
| December 1993 | 375,825.49 | 4.41 | AO |
| June 1998 | –89,441.73 | –3.14 | LS |
| November 1998 | 267,237.35 | 3.42 | AO |
| $\hat{\sigma}_u =$ 103,690 (without outliers) | | $\hat{\sigma}_u =$ 73,462 (with outliers) | |

[a]See note a to Table 1.

From (11), we can predict the subsequent period and we again calculate the one-period-ahead predictions from $T = 131$ (November 1989), as well as the error associated with this prediction. We again consider the calculation of the *MSPE*. Taking into account the new predicted period, the *MSPE* will now be given by

$$(12) \qquad MSPE = \frac{1}{H}\sum_{i=0}^{H-1} e_{T+i}^{2}(1), \quad T = 131,\ldots .$$

The *MSPE* of the last predictions, $T = (228,...,239)$, exceeds that of the first, $T = (131,...,142)$, by some 156 per cent, thereby confirming the progressive deterioration in the degree of forecast accuracy of the previous model.[9] As a

[9]Furthermore, the estimation of the model with information base up to November 1989 has a standard deviation of the residuals of $\hat{\sigma}_u$ = 57,638, and with the information base up to November 1998, this value increases to $\hat{\sigma}_u$ = 91,544. The increase, over time, of $\hat{\sigma}_u$ equals the loss of informative capacity of the predictions made on the basis of the previous model.

consequence, this predictive approach would also indicate the lack of stability in the series when the reference point is December 1989.

However, is the IRPF reform the cause of this break? Put another way, and making use of intervention analysis, if we now model this reform as an event that causes permanent effects to the series, is its estimation significant?

In order to provide a response to this new question, we have introduced into the ARIMA process that is representative of the complete series — expression (2) — two possible formulations of the hypothetical permanent effect associated with the IRPF reform (with a sharp or gradual beginning in December 1989), and have estimated the new expressions. The resultant estimations are

$$(13)\qquad (1-L^{12})y_t = \underset{\substack{10{,}108\\(2.58)}}{\mu} + \underset{\substack{70{,}188\\(2.45)}}{w_1}(1-L^{12})S_t^{t_0} + (1-\underset{\substack{0.52\\(8.91)}}{\Theta_1}L^{12})u_t\,,$$

$$(14)\qquad (1-\underset{\substack{-0.73\\(-3.06)}}{\delta}L)(1-L^{12})yt = \underset{\substack{15{,}573\\(2.15)}}{\mu} + \underset{\substack{155{,}170\\(3.31)}}{w_1}(1-L^{12})S_t^{t_0} + (1-\delta L)(1-\underset{\substack{0.52\\(8.88)}}{\Theta_1}L^{12})u_t\,.$$

In both estimations, the permanent effect is positive and significant (coefficient $w_1$). Note that whilst in expression (13) the additive effect of the reform on tax collection has been modelled assuming that the reform carried with it a permanent tax collection increase as from December 1989, quantified at $w_1$ = 70,188, expression (14) models a tax collection impact that in December 1989 is quantified at $w_1$ = 155,170, but which suffers oscillations and finally stabilises at a value of $w^* = w_1[1/(1-\delta)]$ = 89,528.

On the basis of these last results, everything would appear to indicate that we have found sufficient evidence to support the hypothesis that it was the IRPF reform begun in 1988 that possibly motivated the structural break in the IRPF tax collection series. We can see that tax collection behaviour is different prior and subsequent to the beginning of this reform, as demonstrated by the time-series model identification procedure, and that the presence of this distinct pattern of behaviour is explained by the positive impact that this reform had on the tax collection figures.

## IV. CONCLUDING REMARKS

The aim of this paper has been to detect the hypothetical permanent effects on IRPF compliance that can be associated with the Spanish tax amnesty of 1991. The very limited set of information with which time-series methods operate, together with the adequacy of some of their procedures and instruments for this aim, justify their application.

The first results obtained by way of the Box–Jenkins methodology, which confirm the absence of structural permanence in the tax collection series, are indicative of the fact that some event took place during the period analysed that contributed towards altering the behaviour of the series, thereby introducing permanent effects on the tax collection variable.

In turn, the intervention analysis shows that this event is not related to the granting of the 1991 tax amnesty, and it does not attribute any tax collection effect to the regularisation in either the short or the long term.

The identification of the genuine source of the break receives support from the empirical evidence provided by the Bai and Perron test and from the detection of the outliers in the stochastic data-generating process of tax collection. Both methods coincide in indicating the 1988–91 IRPF reform itself as a possible event that caused the permanent shocks. The repetition of the time-series modelling exercise and of the intervention analysis confirms that this reform introduced a positive and long-lasting impulse to the tax collection series — to which another series of economic and administrative factors could also have contributed — which gave rise to the lack of stability observed in the modelling.

Our results with respect to the effects of the 1991 regularisation procedure on long-term compliance are in line with those obtained by Alm and Beck (1993), in that they provide further evidence of the lack of impact of this type of measure on tax collection figures.

The neutrality of the 1991 amnesty in terms of its effects could have two possible explanations: first, that the 'amnesty effect' was not significant because the final figures that emerged from the regularisation process were not themselves significant, in such a way that a change in the habitual behaviour of taxpayers could not be expected; secondly, that in practice there was a neutralisation of the positive and negative effects associated with this process. If this is indeed the final explanation, then isolating and evaluating the separate effects of the amnesty and the support measures would not appear to be an easy task.

It could perhaps be suggested that the virtue of this regularisation procedure ultimately lay in its ability to bring to an end a refuge for undeclared income that was actually being sustained by government itself. To this, we should add extra-fiscal reasons, in that the low return attached to Special Public Debt undoubtedly contributed to a saving in interest payments and, as a result, to a fall in the State's non-financial cash deficit.

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