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STATE INCOME TAX AMNESTIES I: CAUSES

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Abstract

The purpose of this paper is to analyze empirically for the years 1980-88 the factors which led states with state income taxes to run tax amnesty programs. We find a principal factor to be the level of IRS auditing; in particular, we find that states have tended to "free-ride" on the IRS—if the IRS is active in a state, then that state is less likely to run a tax amnesty program. Indeed, our estimates indicate that had the IRS audit rate remained constant during the 1980-88 period (instead of falling by almost one-half), then the cummulative probability that an average state would have a tax amnesty by 1988 would have fallen by almost one-half compared to its actual level.

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1 Introduction

Between December 1981 and December 1989, 27 states with state income taxes offered some form of tax amnesty program that included income taxes. [1] Another state (Virginia) had a tax amnesty program from February 1, 1990 through March 31, 1990. Two states (Illinois and Louisiana) have each conducted two tax amnesty programs. The dates for state tax amnesty programs which have included income taxes, [2] their gross revenues and coverage characteristics are shown in Table 1.

Gross revenues from state tax amnesty programs have ranged from lows of $150,000 and $280,000 for North Dakota and Louisiana, respectively, to highs of $182 million and $401 million for New Jersey and New York, respectively. All of the programs included nonfilers but varied concerning whether taxpayers with delinquent accounts or taxpayers who filed returns but underreported their taxes were eligible. The earliest five state tax amnesty programs covered nonfilers only.

Despite the fact that 28 of the 40 states with nontrivial state income taxes have now run tax amnesty programs, the efficacy of such programs remains uncertain. Controversy also surrounds proposals for a federal income tax amnesty (Leonard and Zeckhauser, 1987). One reason for these controversies is that no one has yet analyzed empirically, even in modestly sophisticated ways, the factors that have caused states to run tax amnesty programs, the reasons why taxpayers participate in them, or the future consequences of the programs either in terms of the number of returns filed or total revenue collected.

There is a corresponding dearth of theoretical analysis, although some work has occurred, notably by Alm and Beck (1986), Andreoni (1988) and Malik and Schwab (1988), dealing with the participation in amnesties by underreporters. All of these authors consider the revenue impacts of amnesties, but none focus their theoretical analysis on identifying the factors which might increase the likelihood that a given state would decide to initiate a tax amnesty program.

Alm and Beck analyze empirically (for the calendar years 1982–85) those aspects of state tax amnesty programs which tend to affect gross amnesty revenues. Their analysis identifies two features as important: the participation of known delinquents,

[1] Alaska, Florida, Nevada, South Dakota, Texas, Washington and Wyoming have no state income tax. Connecticut taxes only interest, dividends and capital gains, while Tennessee and New Hampshire tax only interest and dividends.

[2] Texas and Florida also have run tax amnesty programs, but these states do not have income taxes.

i.e., the inclusion of accounts receivable, and the coupling of increased enforcement efforts to the amnesty program.

The purpose of this paper is to analyze empirically for calendar and fiscal years 1980–88 the factors which led states with state income taxes to run tax amnesty programs.[3] We find a principal factor to be the level of IRS auditing; in particular, we find that states have tended to "free–ride" on the IRS—if the IRS is active in a state, then that state is less likely to run a tax amnesty program. Indeed, our estimates indicate that had the IRS audit rate remained constant during the 1980–88 period (instead of falling by almost one-half), then the cumulative probability that an average state would have a tax amnesty by 1988 would have fallen by almost one–half compared to its actual level.

In Section 2 we review the relevant literature, concentrating on the characteristics of state tax amnesty participants and the perceived benefits of state tax amnesties to states. We next develop a discrete–time–duration model of amnesties in Section 3. In Section 4, drawing upon the discussion in Section 2, we specify and estimate our duration model and in Section 5 we discuss the results obtained from it. We conclude in Section 6 with some brief speculations on the policy implications of our results.

2 The Objectives of Amnesties and Characteristics of Amnesty Participants

In the most comprehensive published review of state tax amnesty programs, Ross identifies four goals of tax amnesty programs: "(1) place additional taxpayers on the tax rolls and improve future compliance; (2) speed up collections and produce a short–term revenue windfall; (3) create data concerning patterns of taxpayer noncompliance and identify specific areas where enforcement reforms are necessary; and (4) act as a lightening rod to attract public attention around programs to increase enforcement." (1986, p. 152). Ross reports that "[m]ost state officials said that the primary goals of ... amnesty programs were prospective: to get taxpayers back on the tax rolls, and to improve future compliance," the latter in part by "publiciz[ing] increased enforcement mechanisms that were to go into effect immediately after their amnesty periods."

Leonard and Zeckhauser (1987) identify four benefits and three costs of tax amnesties. The benefits are: (1) reduction in the guilt of evaders (they treat guilt as a deadweight loss); (2) increase in revenue from voluntary tax payments; (3) addition of former delinquents to the tax rolls; and (4) smoothing the transition to a regime of stricter tax law enforcement. The costs are: (1) increased feelings by honest taxpayers that the tax system is unfair; (2) encouragement of future noncompliance (due to the anticipation of future amnesties) and (3) reduction in the sense that tax evasion is wrong.

[3]Such an analysis is not only logically antecedent to the investigation of the consequences of running state tax amnesties, it is required of any such investigation. This is because a proper empirical analysis of the gross revenue due and attributable to state amnesties must account for the self–selection of states in choosing to initiate an amnesty program.

Fisher, Goddeeris and Young, in a review of six large state tax amnesty programs (California, Illinois, Iowa, Massachusetts, Michigan and New York), identify three primary characteristics of tax amnesty programs: (1) "The individual income tax accounts for the majority of tax evasion cases resolved through amnesty programs;" (2) "most of the ... cases ... involve small payments, often less than $100;" and (3) "not having filed (rather than having filed incorrectly) is the common form of evasion among amnesty participants." (1989, p. 18).

Finally, as noted in the introduction to this paper, Alm and Beck (1986) analyze empirically the effects of various features of state tax amnesty programs on the direct revenue attributed to those programs, identifying the participation of known delinquents and greater funding for post–amnesty enforcement as the two key features.[4]

Based on these studies, states seem to have initiated tax amnesty programs to accomplish the dual goals of increasing revenues and decreasing noncompliance. A desire for more revenue is straightforward and needs little further discussion. In terms of specification of an empirical model, it leads directly to the consideration of variables related to a state's fiscal "health"; e.g., state tax revenue, federal subsidies, and long-term debt—and to variables related to the potential yield from an amnesty—per capita income, unemployment, and, again, state tax revenue. On the other hand, the desire by states for decreased noncompliance merits some further discussion, especially in terms of its connection to federal enforcement efforts.

There is a direct linkage between the activities of state and federal tax agencies. Congress and state legislatures have explicitly provided for exchanges of otherwise confidential tax return and other tax information between the states and the IRS "to increase tax revenues and taxpayer compliance and reduce duplicate resource expenditures."[5] Agreements on the exchange of tax information also explicitly provide that state tax enforcement agencies and the IRS "will develop cooperative return selection and examination programs" to avoid duplicative efforts.[6] The IRS and cooperative states now routinely, for example, synchronize certain audit decisions. Currently 49 states and the District of Columbia have agreements on the coordination of tax information and audits with the IRS.

Many states, however, rely almost exclusively on the IRS for enforcement of state income taxes. As Snavely summarized as recently as 1988, "it is likely that in most states the bulk of audits come from returns in which information is obviously missing or false claims are evident and from cases in which federal claims have been made against a state's citizens" (Snavely, 1988, p. 908). Indeed, Kansas and Pennsylvania, for example, do virtually no auditing at all while Colorado, Hawaii, Louisiana, Nebraska, North Dakota, Ohio, Oklahoma, Virginia, and West Virginia depend almost entirely on information provided by the IRS in conducting state income tax audits (Dubin,

[4]Leonard and Zeckhauser (1987) also consider the indirect consequences of tax amnesties in terms of the effects of such programs on revenue growth. However, while they find that states which have had tax amnesties experience greater revenue growth than those that have not, they do not control for any other factors besides the existence of an amnesty program.

[5]Internal Revenue Service Manual, *Disclosure of Information Handbook*, §(33) 00.

[6]Id. at Exhibit (33) 00-1, Section 5.1 (Draft Agreement on Coordination of Tax Administration).

Graetz, and Wilde, 1989a).

But IRS audit rates have fallen dramatically over the last decade from roughly two percent to about one percent (Dubin, Graetz, and Wilde, 1989c). This decline has had significant effects on many state tax enforcement programs. In terms of model specification, the fall in federal audit rates coupled with the general reliance of states on information provided by the IRS suggests the consideration of variables related to the IRS's enforcement activities within a state—in particular, the IRS audit rate.

In the next section, before specifying our empirical models, we describe our econometric approach. In general, we attempt to define the probability that a state which has not initiated an amnesty does so, conditional on its past experiences, its economic characteristics, and related IRS activities. In Section 4 we specify the equations we estimate, discussing in more detail the variables used in those equations.

3 A Discrete–Time Duration Model

In this section we describe the econometric approach that is used to estimate the length of time states wait before initiating a tax amnesty. The econometric model we employ is a discrete–time–duration model with normally distributed hazards.

In each period we classify the states according to their participation status: participating in a tax–amnesty program or not yet participating. We specify the probability that a state which has not initiated an amnesty does initiate such a program as conditional on its past experience, its economic characteristics and other factors which may affect the attractiveness of amnesties.

Let y_{it} denote the participation status of state i in period t:

$$y_{it} = \begin{cases} 0 & \text{if not participating during period t;} \\ 1 & \text{if state initiates an amnesty during period t.} \end{cases}$$

The conditional probability that after $t-1$ periods of nonparticipation the state begins an amnesty program in period t, sometimes called the "escape probability," is given by

$$\begin{aligned} P_{it} &= \Pr(y_{it} = 1 | y_{i1} = \cdots = y_{i,t-1} = 0,\ x_{it}) \\ &= \Phi(\gamma' x_{it}), \end{aligned}$$

where x_{it} is a vector of characteristics thought to be related to the probability of initiating an amnesty and $\Phi(\cdot)$ denotes the cumulative normal distribution function. The probability that a state which has not initiated an amnesty up to period t fails to initiate an amnesty in that period is $1 - P_{it}$.

For each state we calculate the number of years that have transpired without any amnesty program and denote this by T_i. The waiting time T_i takes on a maximum value of nine years in our data and we record any state that has waited longer than nine years as a "censored" observation. Censoring occurs because the panel is of fixed length so that the length of time before a state initiates an amnesty cannot be determined for all states at the end of the sampling period.

Let c_i denote the state's status after $T_i - 1$ periods of nonparticipation:

$$c_i = \begin{cases} 0 & \text{if state has an amnesty} \\ 1 & \text{if state does not have an amnesty.} \end{cases}$$

Observations for which $c_i = 1$ are censored, as the observed duration T_i does not represent a completed spell. The treatment of censoring is of some importance in the modeling of amnesty decisions as fourteen of the 40 states which have nontrivial state income taxes did not initiate an amnesty in the nine years for which we have data.[7]

It follows that the conditional probability of observing a spell of length T_i which is either complete ($c_i = 0$) or censored ($c_i = 1$) is:

$$L_i(T_i, c_i) = P_{i,T_i}^{1-c_i}(1 - P_{i,T_i})^{c_i} \prod_{t=1}^{T_i-1} (1 - P_{it}). \tag{1}$$

From (1) we form the log likelihood for the full sample: $L = \sum_{i=1}^{N} log L_i(T_i, c_i)$. Maximizing L with respect to the unknown parameter γ yields consistent and asymptotically normal estimates; optimization is undertaken using a Newton–Raphson algorithm with optimal step size. [8]

Once the parameters have been estimated it is straightforward to obtain an estimate of the expected amount of time states will spend without participating in amnesty programs and the cumulative probability of initiating an amnesty as a function of the duration of time without an amnesty program. The probability of initiating an amnesty program in period s after $s - 1$ periods without an amnesty program is given by:

$$Q_i(s) = P_{is} \prod_{t=1}^{s-1} (1 - P_{it}), \tag{2}$$

so that the expected length of time state i will wait is:

$$E(T_i | c_i = 0) = \sum_{s=1}^{\infty} s Q_i(s). \tag{3}$$

Before estimating the escape probability for amnesties, P_{it}, it is useful to consider the importance of censoring on the estimated duration time. Consider the form of equation (1) when the escape probability P_{it} is a constant P. The sample likelihood becomes:

$$L = \prod_{i=1}^{N} L_i(T_i, c_i) = \prod_{i=1}^{N} P^{1-c_i}(1 - P)^{c_i} \prod_{t=1}^{T_i-1} (1 - P) = \prod_{i=1}^{N} P^{1-c_i}(1 - P)^{T_i+c_i-1}. \tag{4}$$

[7]We estimate our model for the years 1980–88. North Carolina, in fact, ran a tax amnesty in 1989 but we cannot use this information since we do not have the necessary socio-economic factors for 1989.

[8]Estimation is performed within the Statistical Software Tools econometric package. See Dubin and Rivers (1988).

Maximizing L with respect to the unknown probability P gives $\hat{P}_{ML} = (1 - \bar{C})/\bar{D}$ where $\bar{C} = \frac{1}{N}\sum_{i=1}^{N} C_i$ is the percentage of cases which are censored and $\bar{D} = \frac{1}{N}\sum_{i=1}^{N} T_i$ is the average duration of observed waiting time in the sample. Since the expected waiting time is, in this case, the reciprocal of the escape probability, the effect of censoring is to increase the estimate of expected waiting time by the reciprocal of the percentage of cases which are not censored.

In our sample the fraction of censored cases is 14/40. The sample average duration of observed waiting time is $\frac{1}{N}\sum T_i = 257/40$. Thus, the escape probability is estimated to be 0.101 and corresponds to an expected waiting time of 257/26 – roughly ten years. We do not expect the conditional escape probabilities to be constant either across states or in time. The estimated duration models considered below contain the constant probability specification as a special case. As we shall see, these models are rejected at the usual levels of statistical significance in favor of other specifications that allow covariates.[9]

4 Specification and Estimation

We begin with the specification of the escape probability. The escape probability is assumed to be a function of the unemployment rate, personal income per capita, the percentage change in real income tax collections from the previous year and the rate of IRS auditing of individual tax returns. The unemployment rate and per capita income are related to both revenue and compliance, the percentage change in real income tax collection is related primarily to revenue, and the IRS audit rate is related primarily to compliance.

The potential relationship between the unemployment rate and state income tax amnesties is complex. First, states with higher unemployment rates may have unsound economies and thus an amnesty in such a state would produce less revenue. This suggests a negative relationship between the unemployment rate and the likelihood of an amnesty. Second, as indicated in Section 2 above, all amnesty programs include nonfilers, many exclusively so, and the number of nonfilers should increase as the unemployment rate rises. This suggests a positive relationship between the unemployment rate and the likelihood of an amnesty. Third, if unemployment is associated with the so-called "underground economy" generally, then states with higher unemployment rates should have greater noncompliance problems. Again, this suggests a positive relationship between the unemployment rate and the likelihood of an amnesty. On net, since the presence of per capita income in our model should mitigate the first effect to some extent, we expect a positive relationship between the unemployment rate

[9] An alternative specification for the cumulative rate of adoption of amnesties would be the product life-cycle model (Bass, 1969). The product life-cycle model emphasizes the "fad" or "band-wagon" aspects of the demand for a good of which repeat purchase is uncommon. A basic assumption in these models is that the rate of adoption in a given period is proportional to the cumulative activity up to that period. It is therefore noteworthy that our empirical approach, which is based on disaggregate data, also rejects specifications of the adoption rate which are simple functions of elapsed time.

and the likelihood of an amnesty.

Per capita income potentially has a complex relationship with state income tax amnesties as well. First, there is a direct yield effect—states with higher per capita income can expect greater revenue generally from a tax amnesty. On the other hand, such states may be less likely to be experiencing fiscal stress and thus would be less likely to run an amnesty program. Second, states with higher per capita income may have more taxpayers above the minimum income level required for filing, and thus might expect more nonfilers to participate in an amnesty. Third, higher income generally is associated with increased opportunities to evade taxes, and thus states with higher per capita income may have more serious compliance problems. All of these factors except the fiscal stress effect suggest a positive relationship between per capita income and the likelihood of an amnesty.

The percentage change in state income tax collections provides an ideal test of the yield hypothesis versus the fiscal stress hypothesis; i.e., between the hypothesis that states with a solid revenue base are more likely to have an amnesty and the hypothesis that states experiencing fiscal stress are more likely to have an amnesty. A positive relationship between the percentage change in state income tax collections and the likelihood of an amnesty supports the yield hypothesis, while a negative relationship supports the fiscal stress hypothesis.

Finally, we include the federal audit rate of individual returns. This variable simultaneously captures noncompliance and free–riding. On one hand, states with higher federal audit rates might have compliance problems that are known to the IRS. This perception could lead a state to initiate its own amnesty program to combat the noncompliance. On the other hand, states where auditing is high benefit from the presence of the federal government with respect to their attempts to enforce tax compliance. Such states may view federal enforcement efforts as a cheap alternative to their own enforcement efforts and may thus eschew amnesty programs on the grounds that they are both costly and unnecessary.

Among these factors, the role of audits is perhaps the most complicated. An unobserved effect such as greater noncompliance in a state may increase or decrease the state's propensity to initiate an amnesty program. Concurrently, the IRS is likely to respond to the compliance problem and increase its audit rate. To determine which effect (free–riding or directly combatting noncompliance) dominates a state's decision to initiate participation in an amnesty program requires that the simultaneous determination of audit rates and durations be explicitly recognized in the econometric analysis. In this regard, let

$$y^*_{1it} = y'_{2it}\delta_1 + x'_{1it}\beta_1 + \mu_{1it} \tag{5}$$

denote the latent variable for the escape event (initiate an amnesty) in period t for state i. This unobserved measure is a function of state specific exogenous characteristics (x_{1it}) and the potentially endogenous audit level (y_{2it}). We observe the outcome:

$$y_{1it} = \begin{cases} 1 & \text{if } y^*_{1it} > 0 \quad \text{(initiate an amnesty in period t)} \\ 0 & \text{otherwise} \quad \text{(waits at least another period).} \end{cases} \tag{6}$$

The audit rate is specified by a reduced form

$$y'_{2it} = x'_{2it}\Pi_2 + \nu'_{2it} \tag{7}$$

where x_{2it} is a vector of state and time varying characteristics which affect the level of auditing. The econometric difficulty is that the probit specification for the conditional hazard as represented by equations (1) and (2) contains the potentially endogenous regressor y_{2it}. While several solutions to this problem have been proposed (see e.g., Rivers and Vuong, 1988 for a summary), we follow the approach outlined by Smith and Blundell (1986).

Smith and Blundell demonstrate that consistent estimates of the parameters δ_1 and β_1 in the structural equation may be obtained if the estimated residuals from equation (7) are included as additional regressors in the probit function. To obtain consistent estimates we therefore specify

$$P_{it} = \Phi(y'_{2it}\gamma_1 + x'_{1it}\beta_1 + \hat{\nu}'_{2it}\alpha) \tag{8}$$

where $\hat{\nu}'_{2it} = y'_{2it} - x'_{2it}\hat{\Pi}_2$.

As demonstrated by Rivers and Vuong (1988), this procedure does not attain the Cramer–Rao lower bound that is achievable using a limited information maximum likelihood estimator, but it has the advantage of computational simplicity and provides a direct test for the endogeneity of y'_{2it} in the structural equation. Specifically, the endogeneity of y'_{2it} in the probit hazard (4) is equivalent to the rejection of the null hypothesis that the coefficients α are zero.[10]

The reduced form estimates for the audit equation are adapted from earlier work of our own (Dubin, Graetz, and Wilde, 1989b). We take the audit individual rate to be a function of education (ED), age (PER45), unemployment (UR), per capita real income (PICAP), percent employed in manufacturing (PMAN), percent employed in service (PSERV), average state tax rates (STAXR), IRS budget levels (BPR), and time (TIME).[11]

The definitions and sources of these variables (as well as the others used in the duration model) are given in Table 2. We provide mean values of the variables for each of the nine years (1980–1988) for which we have data in Table 3. The means are taken across the 50 states and tabulated by year.

[10] Another method to estimate the parameters of the duration model, would be to substitute the reduced–form for y'_{2it} into the equation for y^*_{1it}. This method is computationally more exacting and does not provide any simple test for the endogeneity for y'_{2it}. A procedure equivalent to that employed by Smith and Blundell (1986) or Lee (1981), (jointly termed the instrumental variables probit (IVP) method) is to substitute the predicted value $\hat{y}'_{2it}$ rather than the residuals $\hat{\nu}'_{2it}$ into the equations for y^*_{1i}. This method requires that the coefficients of $\hat{y}'_{2it}$ and y'_{2it} be summed to obtain the consistent estimate for δ_1.

[11] For a detailed discussion of this specification see Dubin, Graetz, and Wilde (1989b).

5 Results

Table 4 presents the estimated reduced form equation for the audit rates, which has a corrected R–squared of 0.75.[12] The audit rate increases with education, the percent employed in a service industry and with the budget available to the IRS. It decreases with the percentage of the population between 45 and 60 years of age, personal income per capita, the percentage employed in a manufacturing industry, and with time. While these results are for the most part plausible, (audits generally respond to opportunities to evade yet are constrained by the funds that make them possible), we resist the temptation to interpret these reduced form results as structural effects.[13].

Having reduced form estimates for the audit rate, we can proceed to the estimation of the duration models. In Table 5, we provide sample means for the explanatory variables used in the formulation of the escape probabilities. We should note that this table is based on a subset of the 40 states who have some significant state income tax program (as are the estimated duration models). It is clearly not appropriate to model the length of time elapsed before a state initiates an amnesty for any state that is ruled out *a priori* from conducting such a program.

The duration models are based on 360 observations (40 states over nine years). We have calculated the length of time each states waits before it initiates an amnesty and we record this length on both a fiscal and calendar year basis. We display this information in Table 6. Our working assumption is that durations are measured relative to 1980. We use 1980 as a starting point since the first state in our sample to conduct a tax amnesty was Illinois and this amnesty occurred in calendar year 1981. We record this in the table as a one year duration event.

The estimated duration models are presented in Table 7. The results are given for six models. The first three are based on calendar year lengths and the second three models are based on fiscal year lengths. Within each set we estimate first the constant probability model, the model without correction for endogeneity of audits, and the endogeneity corrected model.

The constant probability models agree with the calculations presented above. The constant escape probability is estimated by Model 4 to be $\Phi(-1.27) = 0.102$. However, this model is rejected by the unrestricted model with covariates. In Model 6 for example the log–likelihood is -75.69. The likelihood–ratio test which compares this value against the -83.446 that attains in the constant probability model has a value of 15.5 which exceed the 95 percent critical value of the chi–squared with 5 degrees of freedom 11.1.

The audit rate affect is not significant in the models which are unadjusted for endogeneity (Models 2 and 5). However, once endogeneity is taken into account the

[12]Following Dubin, Graetz, and Wilde (1989b), we assume that the error term in the reduced form equation contains a state specific random effect which causes dependence in the reduced form residuals. The estimates of the reduced form coefficients use generalized least squares to correct for this correlation. We do this to provide correct standard errors for the estimated effects in the reduced form audit equation. It is worth noting, however, that consistent estimation of the duration model only requires consistent (not necessarily efficient) estimates of the reduced form audit estimated residuals.

[13]Again, for a more complete discussion see Dubin, Graetz, and Wilde (1989b)

audit rate is seen to be negative and significant (t–statistics of approximately 1.9).

We therefore find that the federal audit rate is endogenous in the duration models and that its consistent effect is negative. An increase in the audit rate leads states to wait longer before initiating a tax amnesty (the escape rate in any period declines) and thus the hypothesis of states free–riding is accepted—or at least dominates the behavior of the states vis–a–vis their own intentions regarding noncompliance.

Given that the log–likelihoods in Models 3 and 6 are so similar, we do not discern any differences arising from the two methods of measuring the length of time elapsed before a state initiates an amnesty. In the discussion which follows we use the fiscal based specification given in Model 6 for convenience.

The estimated effects in Model 6 are generally in accord with our expectations. An increase in unemployment increases the likelihood of an amnesty as does an increase in per capita income.

Finally, the fiscal stress variable, PSTAX, which measures the percentage change in state tax revenue, has a coefficient which is only marginally significant and positive. If this result is credited as being significant then we find that states for which real state tax collections are increasing are more likely to initiate amnesties, thus supporting the yield hypothesis as opposed to the fiscal stress hypothesis.

To further check the validity of this last result we attempted to measure fiscal stress in a variety of alternative ways. We investigated specifications which included the ratio of state income tax to other revenue sources, the percentage change in this latter variable over time, the percentage change in nonstate tax revenue, and alternative methods for inflation adjustment. All of these avenues were similar in that we could not find, in our data, any support for a fiscal stress theory of amnesties.[14]

To illustrate how our model fits the actual experience of the states we have calculated the cumulative percentage of states which had adopted an amnesty program at a given point in time and we have compared this to the cumulative escape probability as predicted by our estimated duration model. From equation (2), the cumulative escape probability after l periods is given by

$$\sum_{s=1}^{l} Q_i(s) = P_{i1} + (1 - P_{i1})P_{i2} + (1 - P_{i1})(1 - P_{i2})P_{i3} + \cdots + P_{il}\Pi_{t=1}^{l-1}(1 - P_{it}). \qquad (9)$$

We plot the predicted and actual cumulative adoption rates in Figure 1. Our model fits the actual experience of the states (in aggregate) quite well although it somewhat overpredicts the cumulative adoption rate in the early periods and somewhat underpredicts it in the later periods.

Finally, we can illustrate the important role of audits in contributing to the states rate of adoption of amnesty programs. We use the estimated duration model to calcu-

[14]While the level of delinquent accounts is unlikely to explain the probability that a state runs an amnesty which does not cover accounts receivable, it may help explain the probability that a state runs an amnesty which does cover accounts receivable. Unfortunately, the data needed to test this hypothesis is not available now.

late the cumulative adoption probabilities under an assumed course of events in which the IRS had not lowered the audit rate from its 1980 to 1988 levels. We plot the predicted cumulative adoption probabilities based on the historical course of audits and based on the 1980 audit–level rates in Figure 2. Using the 1980 levels of the audit rate shows that states would not have adopted amnesty programs with anywhere near the fervor we have witnessed. By the end of the sample period (1988), the cumulative adoption percentage would only be about 36% as compared with the actual value of approximately 65%. It would therefore appear that a side effect the federal policy of diminished audit capability resulted in shifting a substantial enforcement burden onto the states, who then found it necessary to substitute their own efforts in place of the federal governments.

6 Conclusion

While state tax amnesties may have resulted in increased rates of revenue growth for those states which ran them (Leonard and Zeckhauser, 1987), it is hard to escape the conclusion that many states initiated tax amnesties as part of a systematic effort to respond to the decade long fall in federal tax enforcement activities. Many, but not all, states coupled tax amnesties with increased post–amnesty enforcement efforts of their own. At the same time, there is no evidence that states that ran amnesty programs were under any "fiscal stress." Indeed, states with high per capita income and high growth rates in real state income tax collections were most likely to run tax amnesties, perhaps because an amnesty in such states was more likely to generate a high yield.

TABLE 1

History of State Tax Amnesties, Revenues, and Coverage

State	Time Period	Gross Revenue	Coverage: Nonfilers	Coverage: Accounts Receivable	Coverage: Underreporters
Illinois (# 1)	12-28-81 to 1-08-82	.089	yes	no	no
Arizona	11-22-82 to 1-20-83	6.0	yes	no	no
Idaho	5-20-83 to 8-30-83	.009	yes	no	no
Missouri	9-1-83 to 10-31-83	.845	yes	no	no
North Dakota	9-01-83 to 11-30-83	.15	yes	no	yes*
Massachusetts	10-17-83 to 1-17-84	85.2	yes	yes	yes
Alabama	1-20-84 to 4-1-84	3.1	yes	no	yes
Kansas	7-01-84 to 9-30-84	.6	yes	no	no
Oklahoma	7-01-84 to 12-31-84	17.0	yes	yes	yes
Minnesota	8-01-84 to 10-31-84	12.1	yes	yes	no
Illinois (# 2)	10-1-84 to 11-30-84	158.6	yes	yes	yes
California	12-10-84 to 3-15-85	197.0	yes	yes	yes
New Mexico	8-15-85 to 11-13-85	13.6	yes	no	yes**
South Carolina	9-1-85 to 11-30-85	8.9	yes	yes	yes

* The North Dakota tax amnesty was only open to persons not "under investigation." In principle underreporters were included, but only nonfilers participated.

** Approximately 95% of gross revenue from the New Mexico tax amnesty came from nonfilers.

State	Time Period	Gross Revenue	Coverage		
			Nonfilers	Accounts Receivable	Underreporters
Wisconsin	9-15-85 to 11-22-85	27.3	yes	yes	yes
Colorado	9-16-85 to 11-15-85	6.4	yes	no	yes
Louisiana (# 1)	10-1-85 to 12-31-85	1.2	yes	no	no
New York	11-01-85 to 1-31-86	401.3	yes	yes	yes
Michigan	5-12-86 to 6-30-86	109.8	yes	yes	yes
Mississippi	9-1-86 to 11-30-86	1.0	yes	no	yes
Iowa	9-02-86 to 10-31-86	35.1	yes	yes	yes
West Virginia	10-01-86 to 12-31-86	10.1	yes	yes	yes
Rhode Island	10-15-86 to 1-12-86	1.9	yes	no	no
Arkansas	9-01-87 to 11-30-87	1.2	yes	no	yes
Maryland	9-01-87 to 11-02-87	34.6	yes	yes	yes
New Jersey	9-10-87 to 12-08-87	182.0	yes	yes	yes
Louisiana (# 2)	10-01-87 to 12-15-87	.28	yes	no	no
Kentucky	9-15-88 to 12-15-88	60.1	yes	no	yes
North Carolina	9-1-89 12-01-89	37.6	yes	yes	yes
Virginia	02-01-90 03-31-90	32.3	yes	yes	yes

TABLE 2

Variable Definitions and Sources[1]

CONSTANT	Constant term
ED	Percentage of the adult population with at least a high school education[2]
PER45	Percentage of the adult population between 45 and 60 years of age
UR	Unemployment rate
PICAP	Personal Income per capita in constant (1972) dollars
PMAN	Percentage of labor force that works in a manufacturing industry
PSERV	Percentage of labor force that works in a service industry
STAXR	Average state income tax rate[3]
BPR	Budget per return—IRS state level budget divided by the total number of returns filed[4]
IAR	Individual Audit Rate—number of individual returns examined divided by the number of individual returns filed[4]
PSTAX	Percentage change in real state income tax collections[3]
IARRES	Estimated residual from reduced form audit equation
TIME	Time trend

Notes: (1) Unless otherwise noted, data is taken from *Statistical Abstract of the U.S.*, 1981–1989.

(2) The percent of the adult population with at least a high school education in 1980 is taken from the U.S. Census. For subsequent years we combine data on the number of high school graduates with other demographic data to construct projections based on the 1980 values.

(3) Total state income taxes are taken from *State Government Tax Collections*, 1980–1988, published by the U.S. Department of Commerce. We divide this variable by total state income, taken from *Statistical Abstracts of the U.S.*, 1981–89, to obtain the average state income tax rate.

(4) Individual returns filed, individual returns examined, IRS state–level budgets, and total returns filed are taken from *Annual Report of the Commissioner of Internal Revenue*, 1980–88.

TABLE 3

Sample Statistics for Variables in Reduced Form Audit Equation[1]

Variable[(2)]	1980	1981	1982	1983	1984	1985	1986	1987	1988
ED	.67	.68	.68	.67	.68	.68	.68	.68	.68
PER45	.27	.26	.26	.25	.25	.25	.25	.25	.25
UR	.068	.073	.093	.093	.073	.071	.069	.063	.063
PICAP	5.14	5.24	5.20	5.28	5.54	5.75	5.68	5.71	5.43
PMAN	.21	.20	.19	.19	.19	.18	.18	.16	.16
PSERV	.19	.20	.21	.21	.21	.22	.22	.21	.21
STAXR	1.61	1.58	1.63	1.70	1.80	1.77	1.75	1.64	1.64
BPR	.0045	.0044	.0044	.0048	.0048	.0046	.0042	.0044	.0048
IAR[(3)]	1.66	1.53	1.40	1.36	1.23	1.15	1.00	.91	.83

Notes: (1) Based on all 50 states.

(2) ED, PER45, UR, PMAN, and PSERV are percentages expressed in decimal form (i.e., .67 is 67%). PICAP and BPR are in thousands of 1972 dollars. STAXR and IAR are straight percentages.

(3) These audit rates exclude some audits conducted at the IRS Regional Service Centers.

TABLE 4

Reduced Form Audit Equation[1,2]

Independent Variable	Coefficient	t-Statistic
CONSTANT	13.479	16.9
ED	0.522	1.7
PER45	–5.994	–4.6
UR	–0.318	–0.5
PICAP	–0.078	–3.0
PMAN	–1.537	–4.4
PSERV	1.174	2.4
STAXR	–0.017	–0.5
TIME	–0.134	–25.1
BPR	152.2	9.9
Number of Observations	450.	
Corrected R^2	0.75	

Notes: (1) Sample is based on the 50 U.S. states over the 9 year period from 1980–1988.
(2) Estimated model is based on a random effect specification of the error term.

TABLE 5

Sample Statistics for Variables in Duration Model[1]

Variable	1980	1981	1982	1983	1984	1985	1986	1987	1988
UR	0.070	0.074	0.097	0.095	0.074	0.072	0.070	0.062	0.062
PICAP	5.03	5.13	5.07	5.15	5.42	5.39	5.59	5.60	5.33
PSTAX	–0.0071	0.033	0.018	0.067	0.229	0.053	0.022	–0.044	–0.040
IAR[(2)]	1.58	1.47	1.34	1.29	1.15	1.09	0.92	0.83	0.77

Notes: (1) Based on a subset of 40 states which have a significant state tax program. States excluded are: Alaska, Connecticut, Florida, Nevada, New Hampshire, South Dakota, Tennessee, Texas, Washington, and Wyoming.

(2) These audit rates exclude some audits conducted at the IRS Regional Service Centers.

TABLE 6

Amnesty Program Participation Waiting Time

State	Calendar Based Duration	Fiscal Based Duration
Alabama	4	4
Alaska	no state income tax	
Arkansas	7	7
Arizona	2	3
California	4	5
Colorado	5	5
Connecticut	trivial state income tax	
Delaware	9	9
Florida	no state income tax	
Georgia	9	9
Hawaii	9	9
Idaho	3	3
Illinois	1	2
Indiana	9	9
Iowa	6	6
Kansas	4	4
Kentucky	8	8
Louisiana	5	6
Maine	9	9
Maryland	7	7
Massachusetts	3	4
Michigan	6	6
Minnesota	4	4
Mississippi	6	6
Missouri	3	3
Montana	9	9
Nebraska	9	9
Nevada	no state income tax	
New Hampshire	trivial state income tax	
New Jersey	7	7
New Mexico	5	5
New York	5	6
North Carolina	9	9
North Dakota	3	3
Ohio	9	9
Oklahoma	4	4
Oregon	9	9
Pennsylvania	9	9
Rhode Island	6	7
South Carolina	5	5
South Dakota	no state income tax	
Tennessee	trivial state income tax	
Texas	no state income tax	
Utah	9	9
Vermont	9	9
Virginia	9	9
Washington	no state income tax	
West Virginia	6	6
Wisconsin	5	5
Wyoming	no state income tax	

TABLE 7

Duration Model for Time Waited Before Initiating an Amnesty[1,2]

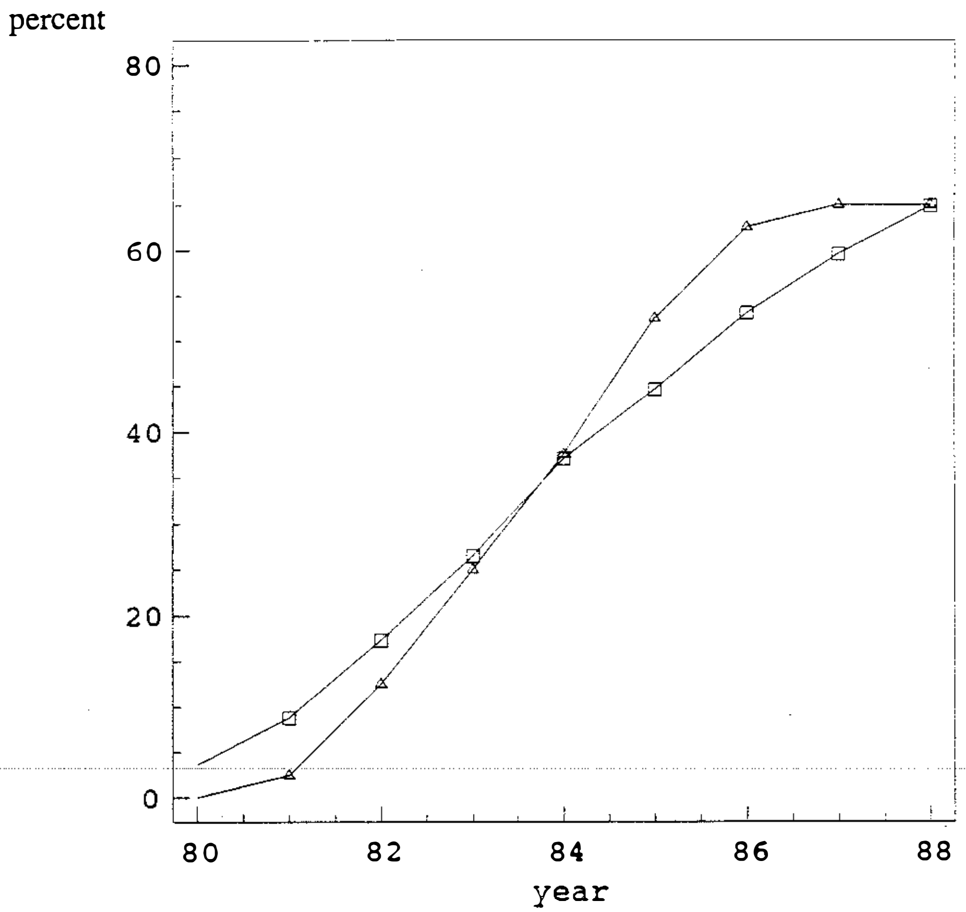
	Calendar Year Duration			Fiscal Year Duration		
Variable	Model 1	Model 2	Model 3	Model 4	Model 5	Model 6
ONE	−1.26 (−11.79)	−3.59 (−3.10)	−3.80 (−3.25)	−1.27 (−12.0)	−3.51 (−3.03)	−3.75 (−3.20)
UR	—	8.73 (1.78)	11.05 (2.15)	—	9.88 (2.02)	12.139 (2.36)
PICAP	—	0.326 (2.00)	0.303 (1.85)	—	0.343 (2.10)	0.325 (1.97)
PSTAX	—	0.966 (1.81)	0.983 (1.68)	—	0.827 (1.84)	0.812 (1.75)
IAR	—	−0.115 (−0.40)	−0.84 (−1.81)	—	−0.34 (−1.151)	−1.075 (−2.26)
IARRES	—	—	1.41 (1.97)	—	—	1.42 (1.98)
Log–Likelihood	−83.446	−77.735	−75.69	−84.204	−78.057	−75.998

Notes: (1) Based on a subset of 40 states which have a significant state tax program. States excluded are: Alaska, Connecticut, Florida, Nevada, New Hampshire, South Dakota, Tennessee, Texas, Washington, and Wyoming.

(2) t–statistics in parenthesis.

FIGURE 1

Actual and Predicted Cumulative Adoption Rates



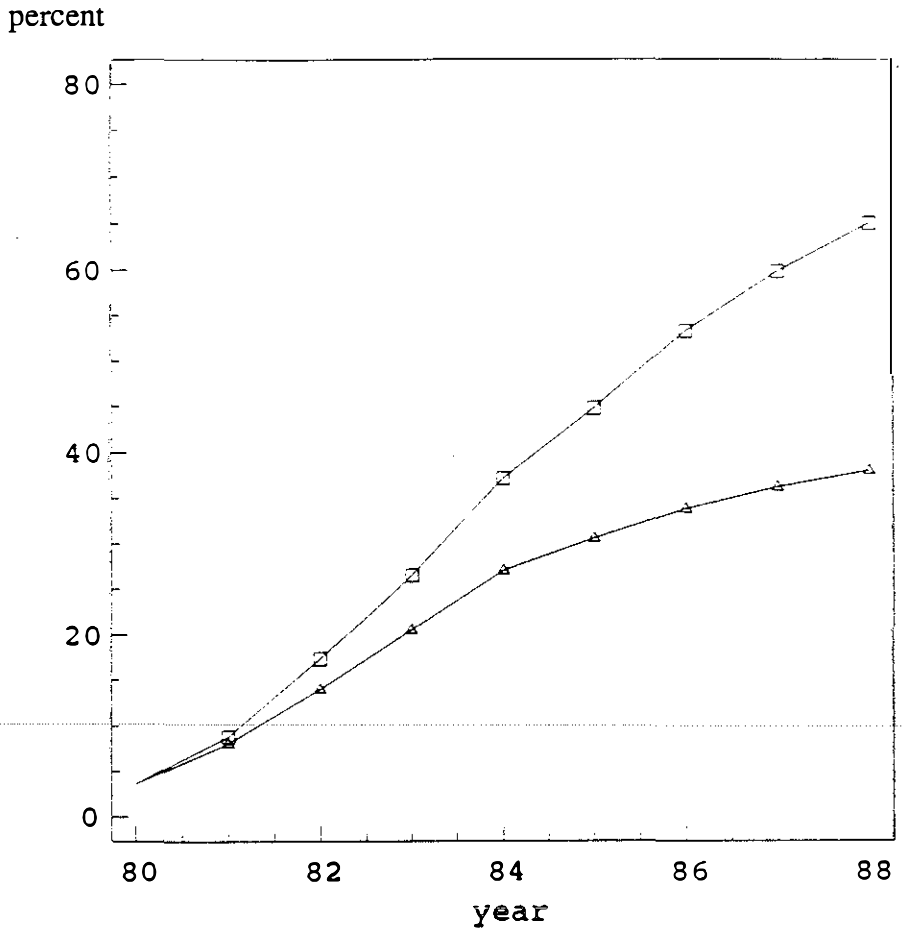
Legend:

□ Predicted Cumulative Adoption Rate from Duration Model

Δ Actual Cumulative Adoption Rate

FIGURE 2

Audit Experiment and Predicted Base Cumulative Adoption Rates



Legend:

□ Predicted Cumulative Adoption Rate from Duration Model

Δ Simulated Cumulative Adoption Rate Had Audit Remained at 1980 Levels

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