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A Model of Welfare Competition with Evidence from AFDC

by

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Abstract

In this paper, we test empirically for strategic behavior among the states using the cash support program Aid to Families with Dependent Children (AFDC). To motivate the empirical work, we adapt Wildasin's [41] model of income redistribution to a model of "interjurisdictional welfare competition." Although welfare competition may be generated from different frameworks, we choose Wildasin's model to illustrate how welfare benefit interdependence is generated in the context of welfare migration. We estimate a "representative reaction function" for AFDC using both cross-sections and pooled cross-section data. After controlling for other determinants of AFDC benefit levels and for spatial error autocorrelation, we find evidence showing competition.

Running head: Welfare Competition.

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1. Introduction

The literature on fiscal competition among local governments has focused mainly on models of tax competition. In a model of tax competition, mobility of the tax base among jurisdictions generates inefficient taxation and provision of local public goods. The leading papers dealing with the small-numbers case, where strategic interaction occurs, are Bucovetsky [12], Mintz and Tulkens [23], and Wildasin [40]. Wildasin [40] notes that local governments may also compete in their expenditure policy, and specifies a model where jurisdictions play a Nash game using the level of expenditures instead of the tax rates as the strategic variable.[1] A direct extension of these models applies to state expenditures on welfare programs. However, in this case, state governments would not be competing to attract tax base but to repel welfare recipients and to keep relatively high income residents. In the tax competition framework, perfect mobility of capital generates competitive behavior among jurisdictions. In the welfare competition model, competitive behavior arises from the mobility of high and low income populations.

The literature on income redistribution has addressed the properties of a decentralized system of income transfers from the rich to the poor. The general result obtained in this literature is that the level of local income transfers is suboptimal due to the presence of externalities arising either from benefit spillovers or from free mobility of individual economic units. Spillover effects arise when the well-being of the poor enters as an argument in the utility function of the nonpoor independently of the location of the poor. Income transfers are too low because spillovers are not taken into account when local governments set the transfer levels. In addition, with population mobility, differences in the level of transfers across local governments induce migration of the poor. Jurisdictions with stronger redistributive policies may also induce the relatively wealthy to move to other less generous jurisdictions (see Stigler, [37], Musgrave [25], and Oates, [28]). As a result, jurisdictions set transfers at levels that are low compared to what they would be if there were no migration.[2] The presence of these externalities generates strategic behavior among local governments in setting the level of income transfers, as in the tax competition framework. Representative authors in this literature include Brown and Oates [8], Orr [30], Pauly [31], and Wildasin [41].

In addition, it has been suggested in the public economics literature that strategic behavior among local governments can also arise from an agency problem between voters and incumbents. Under the "yardstick competition" models, voters make comparisons between jurisdictions to judge the performance of their government officials. This in turn induces incumbents to look to other jurisdiction's policy behavior before deciding any changes in their own policies. Besley and Case [5] and the references there in discuss yardstick competition.

In this paper, we test empirically for strategic behavior among the states using the cash support program Aid to Families with Dependent Children (AFDC). To motivate the empirical work, we adapt Wildasin's [41] model of income redistribution to a model of "interjurisdictional welfare competition." Although welfare competition may be generated from different frameworks, we choose Wildasin's model to illustrate how welfare benefit interdependence is generated in the context of welfare migration. The paper does not attempt to test Wildasin's model nor to determine the "source" of welfare interdependence, but rather to test for it.

The interest in welfare competition has increased because of the recent changes in the welfare system. With welfare reform, the AFDC program has been replaced by the "Temporary Assistance for Needy Families" (TANF) program, which includes among many other changes, the substitution of a block grant in place of matching grants.[3] The switch from the matching grant system to block grants may induce a decline in the welfare benefits, since under the block grant system the marginal cost of welfare spending is entirely borne by the state governments. If federal intervention is needed to raise welfare benefits, as suggested by the analysis, this policy change may be undesirable.

States' autonomy in the determination of AFDC benefit levels made this program a potential candidate for interstate welfare competition. Even though it was jointly financed by states and the federal government under a system of open-ended matching grants, states administered the program, decided eligibility standards, as well as the levels of AFDC payments. Because of this, the AFDC program provides a good test of the welfare competition hypothesis.

Using spatial econometric methods, we specify a reaction function in which the AFDC benefits in a given state are a function of the AFDC benefits in *neighboring* states and own socio-economic characteristics. Our focus is on the slope parameter of the reaction function, which measures the interstate interdependence in the level of AFDC benefits. A non-zero and significant estimate of the slope of the reaction function is interpreted as evidence of AFDC competition. To avoid finding spurious evidence of welfare competition, we test for the significance of this parameter using a generalized method of moments test (GMM) that controls for spatial error correlation.

The results suggest states behave strategically when they set AFDC benefits. The estimated slope parameter of the reaction function is positive and significantly different from zero, indicating that AFDC benefits in any given state are positively affected by the AFDC benefits in neighboring states. This result is robust to different specifications of the reaction function and to the specification of spatial error correlation with individual cross-section estimations as well as with pooled cross-sections and fixed-effects estimations.

Section 2 presents the model of welfare competition. Section 3 develops the empirical specification of the model and discusses the econometric issues. Section 4 examines related empirical studies in the literature. Section 5 discusses the data and presents the results. Section 6 concludes.

2. A Model of Welfare Benefits Competition

2.1. The Model

The model is adapted from Wildasin [41] to follow the formal structure of Bucovetsky's [12] model of tax competition. In this model, the nation is composed of local government units indexed from $i = 1 \ldots I$. The national population is composed of two groups: a group of immobile and identical taxpayers, and a group of mobile non-taxpayers, hereafter welfare recipients. The immobility of taxpayers rules out their migration. They are endowed with all fixed factors of production other than labor, and each welfare recipient is endowed with one unit of labor. The sizes of both groups are nationally fixed. Each jurisdiction produces a numeraire good with a technology represented by $f_i(l_i)$, which is a strictly increasing and concave function of the number of welfare recipients l_i employed in jurisdiction i.

Taxpayers and welfare recipients earn the income generated in the production of the private good. Welfare recipients receive a gross wage given by the marginal productivity of their labor, $w_i = f_i'(l_i)$, and taxpayers earn the remaining income $y_i = f_i(l_i) - f_i'(l_i)l_i$, which corresponds to the gross rents from the fixed factors used in the production of the numeraire good (these rents may include income earned from taxpayers' own labor).

Taxpayers care about the well-being of the welfare recipients located in their respective jurisdictions, and therefore each taxpayer pays a lump-sum tax that finances income transfers to welfare recipients. All welfare recipients within a given jurisdiction receive a transfer of income denoted by b_i, and each taxpayer pays an equal share of the total transfers in his jurisdiction, $b_i l_i / n_i$, where n_i is the fixed number of taxpayers in jurisdiction i.[4] Thus, the arguments of the utility function of a typical taxpayer in jurisdiction i are his own consumption and the consumption of a representative welfare recipient located in his jurisdiction. This utility is represented by $U(y_i, Z_i)$ with both $U_y > 0$ and $U_Z > 0$, and where y_i is taxpayer's net income, and $Z_i = w_i + b_i$ is the net income of welfare recipients. In a non-cooperative game, each jurisdiction chooses its welfare payment to maximize the utility of the taxpayers, taking into account the immigration of welfare recipients, and taking as given the level of welfare payments in the other jurisdictions.

In the strategic tax competition literature, the net return to capital is equalized in equilibrium across jurisdictions due to free mobility of capital. The analogous condition in the welfare competition game is the equalization of the net income of welfare recipients across jurisdictions due to welfare recipients' migration. Thus, across jurisdictions $w_i + b_i = Z$, for some net income Z. Using $f'(l_i) = w_i$ this condition can be written as

$$f'(l_i) + b_i = f'(l_j) + b_j, \quad i \neq j \tag{1}$$

The assumption of a common labor market for welfare recipients guarantees that condition

(1) holds in equilibrium. An increase in welfare benefits by jurisdiction i induces immigration until the income of the welfare recipients is equalized everywhere. The driving force that equalizes net incomes is the change in gross wages caused by the increase in the labor supply due to the welfare-induced migration.

In the context of the AFDC program, this setup can be criticized as unrealistic, on two grounds. First, while the model assumes that welfare recipients earn labor income, reported wage income tends to be negligible for most recipients. However, as noted by Blank [6], observed expenditures levels for many welfare households tend to be unsustainable on welfare benefits alone, suggesting that benefits are supplemented by nonnegligible amounts of unreported income from low paying work. A second criticism concerns the omission of the implicit tax on wage income, which ultimately reduces welfare benefits in a one-for-one fashion.[5] This tax is omitted for simplicity, but the omission may be defensible if recipients' labor income is largely unreported.

Although the framework is subject to the above criticism, its purpose is to generate a simple equilibrium model of welfare migration. A model with similar properties in which welfare recipients do no work could be built on alternate assumptions, such as the presence of moving costs or idiosyncratic attachments to particular locations on the part of welfare recipients (see Smith [34] and Wheaton [39]). However, while such models behave like the present setup, they are more cumbersome to analyze.

The welfare recipients in all jurisdictions must add up to the total population of welfare recipients in the nation. Formally,

$$\sum_{i=1}^{I} l_i = L, \tag{2}$$

where L is the total welfare population. Equations (1) and (2) determine the distribution of welfare recipients across jurisdictions and their common net income Z, conditional on b_i, $i = 1, \ldots n$. Next we look at the comparative statics of the model.

Assuming symmetry, substituting (2) in (1), and differentiating (1) with respect to b_i, the migration response to an increase of the welfare payment in jurisdiction i is given by

$$\partial l_i / \partial b_i = \frac{-(I-1)}{I f_i''(l_i)} > 0, \tag{3a}$$

$$\partial l_j/\partial b_i < 0, \quad j \neq i. \tag{3b}$$

These derivatives indicate that an increase in welfare benefits in jurisdiction i causes welfare recipients to flow from the competing jurisdictions to jurisdiction i. In addition, the net income of welfare recipients in all jurisdictions increases with an increase in the welfare payment in jurisdiction i:

$$\partial Z/\partial b_i > 0. \tag{3c}$$

However, it can be shown that $\partial Z/\partial b_i < 1$. Intuitively, we can see that when there is no migration, the increase in the net income of welfare recipients is equal to the increase in the welfare payment. Thus, $\partial Z/\partial b_i = 1$. However, when there is migration, an increase in the welfare payment causes a decrease in the gross wage due to the in-migration of welfare recipients.[6] Thus migration of welfare recipients decreases the productivity of welfare spending.

Under the assumption that each taxpayer owns an equal fraction of the residual rents accruing to fixed factors in his respective jurisdiction, the net income of a representative taxpayer is given by

$$y_i = \frac{f_i(l_i) - f_i'(l_i)l_i}{n_i} - \frac{b_i l_i}{n_i}, \tag{4}$$

where the first term represents the rents from fixed factors, and the second term corresponds to the share of the welfare cost in his jurisdiction.

2.2. Choice of Welfare Payments

Using (4), the utility of the representative taxpayer in jurisdiction i can be written as

$$U(y_i, Z_i) = U\Big(\frac{f_i(l_i) - f_i'(l_i)l_i}{n_i} - \frac{b_i l_i}{n_i},\ f'(l_i) + b_i\Big). \tag{5}$$

Each jurisdiction chooses its welfare payment b_i to maximize (5), taking into account the migration effect given in (3a), and viewing the welfare payments of the other jurisdictions as fixed. Letting $MRS(y_i, Z) = U_z/U_y$ denote the marginal rate of substitution between private consumption and the consumption of welfare recipients located in jurisdiction i, and using (3a), the first-order condition is given by

$$n_i MRS(y_i, Z) = l_i - \frac{b_i(I-1)}{f''(l_i)}. \tag{6a}$$

This equation has the form of a Samuelson condition for the optimal provision of local public goods. It indicates that in equilibrium, welfare payments in each jurisdiction are set such that the social marginal benefit from increasing the welfare payment equals the marginal cost. The suboptimality of this equilibrium can be seen after we obtain the analogous condition for the no-migration case:

$$n_i MRS(y_i, Z) = l_i. \tag{6b}$$

Because $f''(l_i) < 0$, the RHS of (6a) is bigger than the RHS of (6b), and therefore, the common welfare payment at the Nash equilibrium must be smaller in the migration case than in the no-migration case. This is the key result used in the income redistribution literature to argue that local governments providing welfare will "race to the bottom" in response to welfare-induced migration. The implication is that any decentralized welfare system with no corrections will provide welfare benefit levels that are "too low" from the nation's point of view. It has also been the central aspect of the academic discussion over how a decentralized system of welfare provision should operate and how a system of matching grants compares to a system of block grants under these circumstances.[7]

2.3. Reaction Functions

Note that equation (1) implies that l_i is a function of the welfare payments in jurisdiction i as well as of the welfare payments in the other jurisdictions. This implies that the optimal b_i obtained from the first-order condition is also a function of the welfare payments in the other $I-1$ jurisdictions. Formally,

$$b_i = \Gamma(B_{-i}, X_i), \tag{7}$$

where b_{-i} is the vector of welfare payments of all other jurisdictions in the nation, and X_i corresponds to jurisdiction i's own socio-economic characteristics. This equation constitutes the reaction function of jurisdiction i, and it indicates that the choice of the welfare payment by jurisdiction i depends on the choices of the welfare payments by all other jurisdictions in the nation. It also indicates that jurisdiction i's economic characteristics matter in the determination of welfare benefits. Observe that interdependence of the welfare benefits arises from the fact that the distribution of welfare recipients across jurisdictions depends on the provision of welfare payments in all jurisdictions.

In contrast, in the no-migration case, b_i in (7) is only a function of the socio-economic characteristics of jurisdiction i. When there is no migration, the distribution of welfare recipients is not affected by welfare payments and therefore, each jurisdiction sets its welfare payment without reference to other jurisdictions' benefits. This and some other issues related to (7) are illustrated with the following example of a welfare competition game between two jurisdictions.

Let f be a quadratic function of l, with $f(l_i) = \beta l_i - \alpha l_i^2$, where $\beta,\ \alpha > 0$. In addition assume that the preferences of taxpayers are represented by the function $U(y_i, Z) = y_i + \delta Z - (\gamma/2)Z^2$, where $\gamma,\ \delta > 0$. From (1) $f'(l_1) + b_1 = f'(l_2) + b_2$. Using (2), and substituting $f(l_i)$ in this expression, we obtain the welfare recipients in each jurisdiction. For example, in jurisdiction 1, the equilibrium welfare population is $l_1 = (1/2)L + (b_1 - b_2)/4\alpha$, where L is again the total population of welfare recipients. Note how the population of welfare recipients is a function of the welfare payments in both jurisdictions.

The first-order condition for the maximization of the utility function, i.e. the condition equivalent to equation (6a), is given by[8]

$$n_i MRS(y_i, Z) = \frac{1}{4\alpha}\big(3b_1 - b_2 + 2\alpha L\big). \tag{8}$$

Solving for b_1 we obtain the optimal welfare payment in jurisdiction 1:

$$b_1 = \Omega \ + \ b_2\Big(\frac{1 - 2\alpha\gamma n_1}{3 + 2\alpha\gamma n_1}\Big). \tag{9}$$

where $\Omega = 4\alpha n_1\big(\delta - (\frac{L}{2n_1}) - \beta\gamma + \alpha\gamma L\big)/\big(3 + 2\alpha\gamma n_1\big)$. By contrast, if there is no migration, the optimal choice of the welfare payment in jurisdiction 1 does not depend on the welfare payment in jurisdiction 2.[9]

The term that multiplies b_2 in equation (9) measures the interdependence in the choices of welfare payments among competing jurisdictions, i.e. it is the *slope* of the reaction function. For this particular example, we note that if $0 < \gamma < 1/2\alpha n_1$ the reaction function slopes upward, indicating that jurisdiction 1 increases its welfare payment in response to an increase in b_2. On the other hand, if $\gamma > 1/2\alpha n_1$, the reaction function slopes down, indicating that jurisdiction 1 decreases b_1 in response to an increase in b_2. In addition, if $\gamma = 1/2\alpha n_1$, the reaction function is flat, indicating no interdependence in the choices of welfare benefit levels.

Because it is theoretically possible to obtain either upward or downward sloping reaction functions, the sign of the slope has to be determined empirically by estimating the reaction function of the welfare competition game. We use the state-federal income transfer program Aid to Families with Dependent Children (AFDC) to test for this implication of the welfare competition model. Following (9), we specify a linear version of the reaction function (7),[10] in which state AFDC benefit levels depend on benefit levels in *competing states*, and other relevant socio-economic variables that affect the choices of welfare benefits. The definition of the set of competing states will be discussed in section 3.

Evidence of strategic competition among state governments in setting AFDC benefit levels comes from the slope parameter of the reaction function. If, after controlling for other determinants of AFDC benefit levels, the slope parameter of the reaction function is statistically different from zero, then states respond to changes in the levels of AFDC benefits in other states. In contrast, if the parameter is not statistically different from zero, the conclusion is that the state choices of AFDC benefit levels are independent.

It is important to note that in this paper we do not attempt to test for welfare-induced migration, but for "strategic behavior" among states in setting welfare benefits. In other words, we do not test Wildasin's [41] model, but rather use the model to illustrate how interdependence of welfare benefits is generated in the context of welfare-induced migration. As we noted before, strategic behavior can be generated under different frameworks. On the one hand, strategic behavior may arise if welfare migration is negligible, but if state officials think that it occurs. The fact that some states have tried to impose welfare benefit restrictions for poor migrants suggests that state officials perceive welfare migration and respond to it. On the other hand, interdependence of welfare benefits can also arise from a model of yardstick competition, as noted above. In this framework, voters make comparisons between jurisdictions to judge the performance of their government officials. This in turn induces incumbents to look to other jurisdiction's welfare spending behavior before deciding any changes in their welfare benefits. An interesting application of this model in the context of tax competition is Besley and Case [5].[11]

The following section of the paper presents the empirical specification of equation (7) for AFDC benefits; gives a brief description of this welfare program; and discusses previous empirical work in AFDC benefit interdependencies.

3. Econometric Model

3.1 Specification of the Reaction Function

Equation (7) indicates that local choices of welfare payments are interdependent. This implies that the AFDC benefit level in a given state depends on the benefit levels in competing states. Because our interest is in *testing* for interstate AFDC competition we specify (7) to allow for competition among *all* states and alternatively among *contiguous* states. This enables us to test if AFDC competition is generalized among all states, or if it only occurs among relatively close states. We control for other determinants of AFDC benefits by including socio-economic variables that have been identified in the income redistribution literature as determinants of state AFDC choices.

The empirical version of (7) is given by

$$b_i = \phi \sum_{j=1}^{n} \omega_{ij} b_j + X_i \theta + \epsilon_i, \tag{10}$$

where b_i is the AFDC benefit level in state i; ϕ is a scalar parameter identified as the slope of the reaction function; ω_{ij} is a set of "weights" that aggregates the AFDC benefit levels in competing states; X_i is the vector of socio-economic characteristics of state i; and ϵ_i is the error term, which is assumed to be normally distributed with constant variance and independent across observations. The implications of the assumption of independent errors are discussed further later in this section.

In (10) we relate the AFDC benefit level in each state to a "weighted average" of AFDC benefit levels in competing states, conditional on own-state socio-economic characteristics. Under this specification, the slope parameter of the reaction function measures the response of state i to the change in the *average* level of AFDC in competing states. In this context, (10) is the "representative reaction function" for state i in the welfare competition game. Under this specification, the slope of the reaction function is a *scalar* parameter that applies to the weighted summation of AFDC payments in all other states. A more flexible specification that allows for different slope parameters for each competing state is not feasible because there are insufficient degrees of freedom to identify these and the other parameters of the model.

The weights w_{ij} determine the pattern of interaction among state i and the set of competing states. Because the number of weights increases with the number of observations, it is not possible to estimate them along with the other parameters of the reaction function. This implies that the elements w_{ij} are assumed to be known, and therefore they need to be specified a priori.

We can write (10) for all states in matrix form as

$$B = \phi W B + X\theta + \varepsilon, \tag{11}$$

where B is the vector of AFDC benefits, W is the matrix of weights, X is the matrix of socio-economic characteristics, and ϵ is the vector of errors. The structure of W is such that the off-diagonal elements are different from zero if state i and j are competing states, and equal to zero otherwise. In addition, the diagonal elements of W are set equal to zero, and the rows are standardized to reflect the notion of a weighted average.

We test for welfare competition using weight matrices based on contiguity and distance criteria. We think that state's perception or fear of welfare immigration from other states is stronger the closer these states are. Also, it is more likely that welfare recipients have a stronger migration response to differences in welfare payments in nearby states than to differences in remote states. For example, welfare recipients may have better information about welfare programs in nearby states. Also, migration cost is higher the farther apart states are.

In order to test for welfare competition among contiguous states, we use two *contiguity* matrices. These matrices, denoted W_I, and W_{II}, have off-diagonal elements $w_{ij} = 1/m_i$, and $w_{ij} = (1/d_{ij})/\sum_k(1/d_{ik})$, $j \neq i$, respectively. In this notation, m_i is the number of states contiguous to i; d_{ij} is the mileage distance between the major city in state i and the major city in state j; and the summation is over the states that are contiguous to state i. Also, in both matrices $w_{ij} = 0$, if state i and j do not share borders.

In addition, we test for welfare competition among all states using a distance matrix, denoted W_{III}, with elements $w_{ij} = (1/d_{ij})/\sum_k(1/d_{ik})$, $i \neq j$, where the summation is over all states excluding state i, and d_{ij} is as above. Note that distance matrices aggregate AFDC benefits in competing states using weights that are a decreasing function of distance between a given state and its potential competitors.

3.2 Econometric Issues

In the theoretical model, welfare payments are jointly determined in a Nash equilibrium. Thus, if i and j are competing states, then state i's AFDC benefit level is a function of state j's, and vice versa. This implies that the vector WB on the right-hand side of (11) is endogenous. Due to this endogeneity, ordinary least squares estimates of the parameters of (11) are inconsistent.[12] Assuming that $(I - \phi W)$ is invertible, we solve for the reduced form of the model, which is given by

$$B = (I - \phi W)^{-1} X\theta + (I - \phi W)^{-1}\varepsilon. \tag{12}$$

Note that (12) is non-linear in its parameters. We estimate this model and test for the significance of the slope parameter ϕ using maximum likelihood methods. The maximum likelihood estimates are consistent and efficient under standard regularity conditions, which are in general satisfied if the structure of interaction in the model given by the product of ϕ and W is non-explosive. In general, this condition is guaranteed after imposing restrictions on the values that the parameter ϕ can take. For the weight matrices used in our estimations, this restriction means that ϕ must be less than one in absolute value.

The empirical model specified in (12) is known in the econometrics literature as a spatial lag model. In these models, the weight matrix W is seen as a "spatial lag" that shifts the dependent variable in space (as opposed to time in the time-series context). For specification, testing and estimation of these and other spatial models, see Anselin [1].

A standard test of the significance of the "spatial lag" parameter ϕ is the Anselin [2] test based on the Lagrange Multiplier principle (hereafter denote LM_{ϕ}). However, the power properties of this test are not good when the errors are spatially correlated. Note that in (12) we assumed that the errors are independent across observations. To illustrate the source of the problem, assume for a moment that instead, the errors follow a spatial autocorrelated process described by

$$\varepsilon = \lambda W \varepsilon + \nu, \tag{13}$$

where ν is a well behaved error vector. This error generation process can arise when the error term includes spatially-correlated omitted variables.

Also, suppose that states set AFDC benefit levels with no reference to the benefits of other states. This assumption implies that $\phi = 0$, and therefore (12) becomes

$$B = X\theta + \varepsilon. \tag{14}$$

Equations (13) and (14) together represent a simple regression model with a spatially autoregressive disturbance term. After solving for ε in (13) and substituting its value in (14), this model can be represented by

$$B = \lambda WB + X\theta + WX\delta + \nu, \tag{14a}$$

where $\delta = -\lambda\theta$.

Note, that except for the extra term $WX\delta$, this model has the same "appearance" as the model in (11). The error autoregressive parameter λ appears now as the parameter of the lagged dependent variable WB. The implication of this is that when the true model is given by (13) and (14), a test for the significance of ϕ in (12) may not reject the null hypothesis of a zero value, leading to the false conclusion of welfare competition. Similarly, under these conditions, a test of spatial error dependence in (12) may not reject the null hypothesis of independent errors. These problems arise because under misspecification, tests of $\phi = 0$ or $\lambda = 0$ are not independent of each other.

As discussed in Brueckner and Saavedra [11], one of the approaches to overcome this problem is to use the robust Lagrange multiplier (LM) tests LM_ϕ^* and LM_λ^* developed by Anselin et al. [4]. These tests are robust to local misspecification of the type described above and have been shown to have good power properties even if misspecification does not exist. In addition to the standard test, in this paper we use LM_ϕ^* and LM_λ^* to test for $\phi = 0$ and $\lambda = 0$. If both tests show that ϕ is nonzero and λ is zero, then the probability of finding a "spuriously" significant slope parameter reduces to the size of the tests. Also, if both tests are significant, this suggests that in order to obtain efficient estimates of the parameters of the model, we should estimate the full model, i.e. allowing for spatial error correlation.

In addition, note that if the nonlinear constraint $\lambda\theta = -\delta$, holds, then the model given in (14a) is equivalent to the regression model with autoregressive errors given by (12) and (13) together. The test of this restriction or "common factor hypothesis," as it is known in the literature, indicates whether a spatial error model is appropriate. We also perform this additional test to double check the results from the standard and robust tests.

As has been documented in the spatial econometrics literature, maximum likelihood estimation of the full model suffers from identification problems (Anselin [1], Anselin and Bera [3], and Anselin et. al. [4]).[13] When both LM_ϕ^* and LM_λ^* are significant, in order to test for the significance of ϕ when λ is specified, we estimate the full model using a generalized method of moments procedure (GMM), assuming that the error term follows an autoregressive process like the one described in (13). We use the GMM estimator suggested by Kelejian and Robinson [19].[14] Based on the work of Newey and West [26], we formulate a GMM version of a Lagrange multiplier or score statistic, and use this statistic, hereafter denoted LM_{GMM}, to test for the significance of the slope parameter of the reaction function. The details of the derivation of the GMM estimator and LM_{GMM} statistic are available upon request.[15]

4. Previous Evidence on AFDC Competition

Most of the empirical work related to interstate welfare competition has concentrated on looking for evidence of welfare induced-migration. The empirical findings in this literature are mixed. Some studies report evidence showing migration (Southwick [36], Gramlich and Laren [17], Blank [7], Enchautegui [15], and Meyer [22]), and some other studies report evidence against it (Levine and Zimmerman [21], and Walker [38]).[16]

Few authors have tried to empirically investigate the consequences of welfare competition by estimating reaction functions. Authors who have implicitly estimated reaction functions are Figlio et al. [16], Peterson et al. [32], and Smith [35]. Other studies whose empirical models are related to a model of strategic welfare competition are Shroder [33], Smith [34], and Wheaton [39]. The findings in this literature are more consistent. Most of the studies lead to the conclusion, either directly or indirectly, that there are interstate dependencies in the AFDC program. Only two authors, Shroder [33] and Smith [35], report results that do not favor this hypothesis.

Figlio et al. [16] and Peterson et al. [32] provide evidence that there is strategic competition in setting AFDC benefit levels. The first authors use an instrumental-variables approach to estimate a version in first differences of a reaction function like (12). Instrumental variables estimation allows them to produce results that are robust to spatial error correlation. They find evidence of differential state responses to increases and decreases in their neighbors' benefits. Using maximum likelihood methods, a similar model is estimated by Peterson et al. [32]. To account for contemporaneous error correlation, they include a *temporal lag* of the AFDC benefits among the control variables.[17]

On the other hand, Shroder [33] uses a simultaneous equation approach to study the effects that welfare benefits have on the number of welfare recipients in a state, as well as to explore how the size of the welfare population affects own-state and *neighbors'* benefit levels. Based on the results of this exercise, he finds no evidence of strategic interaction in welfare payments. Smith [35] estimates a single equation like (12). However, the right hand side variable is not a weighted average of contemporaneous benefit levels in other states, but of the *lagged* benefit levels in these states. Therefore, he does not allow for simultaneity in the choice of AFDC benefits.

The present paper complements the empirical literature in welfare competition. Using a rather different econometric approach, we also find evidence that state welfare benefit choices are interdependent.

5. Data and Estimation Results

5.1 Data

We estimate the reaction function using three different cross-section samples of the states in the continental U.S., from years 1985, 1990 and 1995.[18] We also estimate the model using a pooled cross-section of these quinquennial years, including fixed effects to account for time invariant specific characteristics of the states that may affect AFDC benefit levels. As we noted before, the dependent variable corresponds to the maximum AFDC benefit levels for a family of size three across the states. The right hand side variates include variables that measure socio-economic characteristics of the states and AFDC recipients.

The socio-economic variables are per capita state income, the African-American proportion of the population, the proportion of state representatives that are Democrats, average female unemployment, proportion of recipient households that are not married, and the state share of AFDC benefits.[19] These variables have been shown to be important in the determination of welfare benefits, and their use follows the usual practice in the empirical literature. After conditioning on the other variables, per capita state income indicates availability of resources to finance welfare programs, and therefore it is expected to have a positive influence on the level of welfare benefits. The African-American population variable is a proxy for poverty, and its effect is expected to be negative. The idea is that high levels of poverty may induce states to cut benefits to reduce the total cost of a higher welfare caseload. The female unemployment variable proxies for the size of the population of potential welfare recipients. Higher female unemployment may thus increase the number of welfare recipients, inducing states to decrease AFDC payments to avoid providing benefits to more recipients. If this is the case, the coefficient of this variable will be negative. However, previous empirical findings show a positive correlation between this variable and AFDC benefits (Shroder [33], Craig [14], Smith, [35]. A positive coefficient may be interpreted as states increasing benefits to compensate job-losers (Smith [35]).

The party representation of the state legislature is measured by the proportion of State Representatives (i.e., members of the lower house) that are Democrats. This variable measures the political control of the states, and is expected to be positively correlated with the level of welfare benefits. The proportion of recipient households that are not married has been shown in other studies to exert a significant negative effect on AFDC benefits. The explanation of this finding is that states want to avoid giving economic incentives to single parenthood.[20]

The state share of AFDC benefits measures the price effect that influences AFDC benefits through the matching grant system. It is expected that the lower the state share is, the higher is the level of AFDC benefits (see footnote 7). However, as is well known, the state share of welfare benefits is highly correlated with per capita state income. This follows because the federal grant is determined by a matching formula that takes into consideration state per capita income in establishing the state share of welfare costs.[21] For example, the correlation coefficient is 0.91 in the 1990 cross-section sample, and 0.80 in the pooled cross-section sample. This creates difficulties in determining separately the effects of per capita state income and state share on the level of welfare benefits. To see this, assume for a moment that AFDC benefits are determined only by per capita state income (Y), and state share (S), i.e. $AFDC = \alpha Y + \delta S + \epsilon$. Since S depends on income (see footnote 21), when we differentiate AFDC benefits with respect to income, we obtain $dAFDC/dY = \alpha + \delta\ \partial S/\partial Y$, which corresponds to the combined effects of income and the state share.

We identify the effect of the state share variable on AFDC benefits through an indirect method suggested by Craig [14]. We know that the matching rate is bounded from above and from below. States that have relatively high per capita incomes receive a maximum federal grant of 50% of AFDC benefits, states that have relatively low incomes receive a maximum of 83%, and states with intermediate incomes receive a rate that varies between these two extremes. In 1995, eleven states were at the lower boundary, receiving 50% of the AFDC benefits, and no state was at the upper boundary (Mississippi had the highest matching rate of 79%).

For states that are at the lower and upper boundaries, the matching rate is insensitive to changes in income, and therefore, there is no collinearity between the state share and per capita state income for these states.[22] In other words, variation in income in these states does not cause changes in the state share. Craig's [14] method for isolating the effect of the matching rate from the per capita income effect consists of breaking up per capita state income into two variables. One variable is equal to per capita income for states at the lower matching rate boundary, and zero otherwise. The other variable is equal to per capita income for states with a matching rate inside the boundaries, and zero otherwise. Then, assuming that the income effect is equal across states, the coefficient of the first variable measures the effect of per capita state income on AFDC benefits, while the coefficient of the second variable measures the combined effects of the matching rate and per capita income effects for states with a matching rate inside the boundaries. Using these two coefficients and the matching formula, it is possible to determine the effect of the matching rate on AFDC benefits for these states. This effect is obtained by subtracting the value of the coefficient of the first variable from the value of the coefficient of the second variable, and then adjusting the resulting value by the inverse of the derivative of the matching formula with respect to income. In addition, a dummy variable taking values of 0.50 for states at the lower matching rate boundary and zero otherwise (denoted hereafter S_H), is included in the model to capture the effect of the state share for these states.

To illustrate the method, suppose that AFDC benefits are determined only by per capita income and state share, as above. Now, separate the per capita income variable in two variables and call them Y_H and Y_L, respectively. Then, $AFDC = \alpha Y_H + \gamma Y_L + \beta S_H + \epsilon$. Second, differentiate AFDC with respect to income. Note that $\partial S/\partial Y = 0$ for states at the lower matching rate boundary. Therefore, $\hat{\alpha}$ is an estimate of $\partial AFDC/\partial Y$, and $\hat{\gamma}$ is an estimate of $dAFDC/dY$ for intermediate income states (i.e., $\hat{\gamma} = \hat{\alpha} + \hat{\delta}\ (\partial S/\partial Y)$ as explained above). This implies that $\hat{\delta} = \partial AFDC/\partial S = (\hat{\gamma} - \hat{\alpha})/(\partial S/\partial Y)$. Because $\partial S/\partial Y$ depends on state income, as determined by the matching formula, we can evaluate this for an average income state. In this paper, we follow this method to isolate the effects of per capita state income and the matching federal rate. This allows us to analyze the effects of the switch from the matching grant system to a block grant system in the context of welfare competition.

5.2 Estimation Results

Descriptive statistics for all variables in the pooled cross-section sample are presented in table 1. Tables 2, 3 and 4 contain the parameter estimates of the reaction function for the 1985, 1990 and 1995 cross-section estimations, respectively. Tables 5 and 6 contain estimates from the pooled cross-sections and from the fixed-effects estimations. Table 5 presents maximum likelihood estimates and table 6 presents generalized method of moments estimates. The corresponding t-values of the parameter estimates are given in parenthesis. For all cross-sections and for the pooled cross-sections, the reaction function is estimated using the three weight matrices described in section 3. In addition, each table presents test statistics from different error specification tests, and from tests of the significance of the slope parameter of the reaction function ϕ, and the error parameter λ. Probability values ($p-$values) of all tests statistics are given in parenthesis.

5.2.1 Cross-Section Results

Analyzing the role played by the control variables in the cross-section estimations, we observe that, in the 1990 and 1995 cross-sections, the parameter estimates of per capita state income, Afro-American proportion of the population, and the percentage of recipients households that are not married, have expected signs, and are statistically significant at the 95% confidence level under W_{II} and W_{III}, and at the 90% confidence level under the specification with W_I. In addition, the coefficient of the proportion of state representatives that are Democrats, which is always positive, and the coefficient of female unemployment, which is always negative, are significant (at the 95% confidence level) across all specifications of W in the 1990 cross-section estimations, while they are not significantly different from zero in the 1995 cross-section estimations (the political variable, however is significant under W_{III}).

On the other hand, in the 1985 cross-section estimations, only per capita state income and the percent of unmarried have effects that are significantly different from zero at the 95% confidence level under all specifications of the weight matrix. They also have the expected positive and negative signs, respectively. The coefficients of the female unemployment and political variables are not significantly different from zero, while the proportion of Afro-

American population appears negative and significant at the 90% confidence level under the weight matrix W_{III}, and insignificant under the contiguity matrices W_I and W_{II}.

The estimates of the state share variables for intermediate and high income states are not significantly different from zero under any specification of W.[23] They also have the wrong sign in the 1990 and the 1995 cross-section estimations. The cross-section estimations then suggest that the matching rate did not have the expected negative price effect in these three different samples.

Turning to the estimates of the slope parameter of the reaction function ϕ, we observe that they have positive signs in all cross-section estimations and across all specifications of the weight matrix (except in 1985 under W_{III}). In addition, the estimates are consistently highly significant (at the 95% confidence level) under the specification with the contiguity matrix W_I. They also appear to be significantly different from zero at the 90% confidence level under W_{II} in the 1985 and 1990 cross-section estimations, and at the 95% confidence level under W_{II} in the 1995 cross-section estimations. On the other hand, the estimates are not significantly different from zero under the specification with the weight matrix W_{III} in any of the cross-section estimations. It is important to note that each weight matrix represents a different pattern of interaction among a given state and the set of competing states. This implies that the specifications of the reaction function are non-nested, and therefore estimated values of the parameter ϕ can not be compared across W.

Taken together, these results suggest that in these three different periods, state choices of AFDC benefits depended on choices in contiguous states, while they were independent of the AFDC choices in non-contiguous states. In other words, the cross-section estimates provide evidence of AFDC competition among states that share borders.

The hypothesis of normal errors and the hypothesis of constant error variance cannot be rejected at the 5% level of significance for any of the estimated models. Observe that the Kiefer-Salmon and the Breusch-Pagan statistics are not significant. In addition, the Lagrange multiplier test of spatial error dependence is significant only in the 1990 cross-section estimations under the weight matrix W_{II}. In the other two cross-sections, the hypothesis of independent errors cannot be rejected under any of the weighting schemes.

The results from the robust tests are presented at the bottom tables 2, 3 and 4. Observe, that, in the 1990 and 1995 cross-section estimations, the tests values of LM_{ϕ}^{*} indicate that the slope parameter of the reaction function is statistically different from zero at the 5% level, under the specifications with contiguity matrices W_I and W_{II}, while it is not significant under W_{III}. These results are consistent with the findings described above for these two cross-section estimations. However, in the 1985 cross-section estimations, the robust test is not significant at the conventional level of significance under W_I and W_{II}, contradicting the previous results for these two specifications of the reaction function.

The robust test for the significance of λ, which is the parameter of spatial error correlation, indicates the absence of error dependence under all specifications of the weight matrix for the 1985 and 1995 cross-section estimates. However, it indicates spatial error dependence under all specifications of the weight matrix for the 1990 cross-section estimates (the test is significant at the 10% level under W_I and W_{III}, and at the 5% level under W_{II}). This means that the 1990 cross-sample may exhibit spatial error dependence, and therefore, more efficient estimates could be obtained under a full specification of the reaction function, i.e. allowing for spatial lag and spatial error dependence. Note that in this cross-section estimation, the test of the common factor hypothesis indicates that the spatial error model is not appropriate under W_I and W_{II}, which were the specifications where lag dependence as well as spatial error dependence appear to be statistically significant (see bottom of table 3). Taken all together, these tests results suggest that states react to the choices of AFDC benefits of their neighbor states.

Summarizing, cross-section estimations of the reaction function support the hypothesis of welfare competition among states that share borders. During these three different periods, the best response that a given state had to an average decrease in the AFDC benefit levels of its contiguous neighbors was to decrease its own AFDC benefits, as indicated by the positive sign of the estimated slope of the reaction function. Next we turn to the pooled cross-section results.

5.2.2 Pooled Cross-Section Estimations

Two sets of maximum likelihood (ML) estimates are obtained with the pooled cross-section sample. One set of estimates is obtained without including state fixed effects, and another set includes fixed-effects, as seen in Table 5. After pooling the cross-section samples, all the tests for spatial error dependence were highly significant, under all specifications of the weight matrix. Therefore, a second set of non-fixed effects and fixed-effects estimates are obtained with an efficient generalized method of moments (GMM) estimator to account for this spatial error dependence. Table 6 shows these estimates.

The ML estimates from the non-fixed-effects models are given in the first three columns of table 5. All parameter estimates of the control variables have expected signs, except the estimates of the state share variables, which are again positive, but, as in the cross-section estimations, these estimates are not significantly different from zero under any specification of the weight matrix. The other variable that appears insignificant under any specification of W is female unemployment. All the other control variables are significant across W.

The ML estimates with the pooled sample also support the hypothesis of welfare competition among states that share borders. The estimated slope coefficient of the reaction function is significantly different from zero under all the specifications of W. The robust test LM_{ϕ}^{*} is significant at the 5% level under the specifications of the reaction function with the two contiguity matrices, while is not significant under the W_{III} specification. The estimates are again positive, which indicates that the reaction function is upward sloping.

As in the cross-section estimations, we do not reject the hypotheses of normal and constant variance error across all specifications of W. The LM_{λ} and the LM_{λ}^{*} tests of spatial error dependence are significant at the 5% level under the weight matrix W_{III}, and are not significant under the contiguity matrices. However, the common factor hypothesis rejects the spatial error model across all specifications of W.

In short, the pooled cross-section estimations corroborate the statistical evidence of AFDC competition found in the individual cross-section samples.

Similar results are obtained in the fixed-effects model.[24] However, the slope parameter of the reaction function appears to be significant under *all* specifications of the weight matrix, including W_{III}, suggesting that states compete in setting AFDC benefits looking not only at their immediate neighbors. Observe that the estimated slope coefficient is positive and highly significant under the specification with the matrix W_{III}.

After including the state fixed effects, most of the control variables appear to be insignificant in the determination of AFDC benefits. Only the proportion of Afro-American population is significant at the 5% confidence level, across W. The other variables, including the per capita state income, are not significant even at the 90% confidence level. In addition, the estimates of the per capita state income variable have an unexpected negative sign under the contiguity matrices W_I and W_{II}.[25] The sign is again positive under the matrix W_{III}. However, these estimates are not significantly different from zero. Even though the state share variables have the expected negative sign, in contrast to the positive estimates obtained in the cross-sections and in the pooled cross-section models, they are not significant under any specification of the reaction function.

The hypothesis of AFDC competition is robust to the inclusion of fixed-effects. The estimates of ϕ are significantly different from zero under all specifications of W. In fact, the fixed-effects estimates suggest that states choices of AFDC benefits are not only influenced by the AFDC choices of their immediate neighbors, but by the choices of other states, as suggested by the significance of ϕ under the matrix W_{III}, which was not significant in the cross-section and pooled cross-section estimations. However, most of the control variables do not play any role in explaining the AFDC benefits, in contrast to the significant roles found in the non-fixed effects model. Even though some of the explanatory variables, mainly the state share, have been found insignificant and with the wrong sign in previous empirical research, the non-significant per capita state income is rather unexpected.[26]

Turning to the GMM estimates, observe that they also corroborate the hypothesis of AFDC competition among the states. As with the maximum likelihood estimates, the GMM estimates suggests that the phenomenon of AFDC competition is generalized among all states, and not only among states that share borders. The non-fixed-effects as well as the fixed-effects estimates of the slope parameter are significant at the 95% confidence level, under all specifications of the weight matrix.

Observe that the GMM and ML estimates obtained in the non-fixed-effects models are similar. As the ML, the GMM estimates of the right hand side variables (excluding the state share and the female unemployment variables) have the expected signs and appear significantly different from zero. The main difference among these estimates occurs in the fixed-effects models. The GMM estimates of the parameter ϕ obtained in these models are greater than one under the specifications with the contiguity weight matrix W_I and W_{III} (this is also the case for the non-fixed effects model using W_{III}).[27] However, as in the maximum likelihood estimation, only the proportion of Afro-American population turned out to be significant at the 95% confidence level. The other variables, including the per capita state income, which has the expected positive sign, are not significant at conventional significance levels.

Finally, the estimated coefficient of spatial error correlation is significant under all specifications of the weighting matrix in the pooled cross-section as well as in the fixed- effects models. After accounting for it, the main conclusion from these results is that the welfare competition hypothesis is robust to the specification of spatial error dependence.

6. Conclusion

We find evidence that state choices of AFDC benefits are interdependent. This result is robust to the specification of state fixed effects and to the specification of spatial error correlation. A Lagrange multiplier test that is robust to local misspecification of spatial error autocorrelation, as well as a GMM statistic in the full model, suggest a positive and significant slope of the reaction function, supporting the hypothesis that states are competing when setting AFDC benefits. In addition, the estimated coefficient of the slope parameter of the reaction function is significant under different specifications of the weighting scheme. The positive estimates indicate that the reaction function is upward sloping, and therefore, during the periods covered by our samples, the best response that a given state had to an average decrease in the AFDC benefit levels in the other states, was to decrease its own benefits.

The estimates of the state share variables are not significantly different from zero in any of the estimated models, and under any specifications of the reaction function. This finding is also reported by Shroder (33), who finds an insignificant state share effect after including fixed effects. Craig (14) also finds an insignificant coefficient at the conventional 95% confidence level (but significant at the 90% confidence level). This result of a zero price elasticity of AFDC benefits suggests that states have been not responsive to the matching grant system, and therefore, that the switch from this system to the block grant system will not have the expected negative effect on the AFDC benefit levels. If the matching aid has been diverted to other state programs, as has been suggested in the literature (see Craig (1995)), there is no reason to think that the block grants will play a different role.

The results of interdependent choices of AFDC benefits cannot be taken as "proof" that state governments engage in welfare competition as suggested in the theoretical model presented in section 2. As discussed in the paper, our empirical findings could also arise from a model of yardstick competition, or from a framework in which welfare migration does not occur but is just "perceived" by state officials. Further research is needed to determine what is the empirical likelihood of competing theoretical models.

Table 1.
Descriptive Statistics
Cross-Section Samples

Variables	**Means**			**Std. Devs.**		
	1985	1990	1995	1980	1990	1995
AFDC Benefits for a Family of Size 3*	324.8	369.6	381.8	112.8	134.8	135.4
Per Capita Income*	13300	17603	22091	2013	2835	3334
State Share Interm. Inc. States	0.37	0.35	0.35	0.07	0.08	0.07
Population Proportion Afro-American	0.10	0.10	0.10	0.09	0.09	0.09
Prop. Dem. State Representative	0.58	0.59	0.51	0.18	0.15	0.17
Female Unemployment	7.3	5.39	5.16	1.9	1.14	1.10
Pct. Recip. Unmarried Parent	51.8	54.7	56.7	11.3	11.0	9.54

* At current prices.

Table 2.
Cross-Section ML Estimates For 1985

Dependent Variable: Ln AFDC Benefits 1985 for a family of size 3.	Maximum Likelihood Coefficient Estimates[1]		
	Weight Matrices		
Explanatory Variables:	W_I	W_{II}	W_{III}
W AFDC	0.28 (2.13)	0.21* (1.67)	-0.03 (-0.10)
Constant	-6.54 (-2.13)	-6.56 (-1.86)	-7.73 (-1.79)
Ln Per Capita Income	1.95 (2.82)	2.03 (2.89)	2.38 (3.29)
Ln State Share Interm. Inc. States[2]	-0.31 (-0.89)	-0.32 (-0.92)	-0.40 (-1.36)
Dummy State Share High Inc. States	-11.8 (-0.90)	-12.4 (-0.92)	-15.5 (-1.12)
Ln Prop. Afro-American	-0.06 (-1.29)	-0.06 (-1.34)	-0.08* (-1.86)
Ln Prop. Dem. State Representative	0.01 (0.10)	0.01 (0.10)	0.00 (0.07)
Ln Female Unemployment	-0.07 (-0.41)	-0.08 (-0.49)	-0.12 (-0.70)
Ln Pct. Recip. Unmarried Parent	-0.50 (-2.69)	-0.54 (-2.84)	-0.60 (-3.11)
Tests of Error Specification[3]			
LM Breusch-Pagan Heterosked.	5.94 (0.55)	6.38 (0.50)	8.51 (0.29)
Kiefer-Salmon Normality Test	1.09 (0.58)	1.05 (0.59)	2.45 (0.29)
LM Test of Spat. Error Depen.	0.01 (0.92)	0.02 (0.87)	0.14 (0.71)
Robust Tests of Spat. Dep.			
LM Test of Spat. Lag Depen.	2.49 (0.11)	1.64 (0.20)	0.03 (0.87)
LM Test of Spat. Error Depen.	0.24 (0.63)	0.24 (0.63)	0.01 (0.75)
Common Factor Test			
LR Test on Common Factor Hyp.	7.20 (0.41)	8.53 (0.29)	9.70 (0.21)

[1] Asymptotic standard normal statistics are given in parenthesis.
[2] The corresponding estimates of the per capita income coefficient for states with matching rate inside the boundaries are: 1.33 (4.10), 1.39 (4.20), 1.57 (4.65), respectively.
[3] *p-values* are given in parenthesis.
* Significant at a 90% confidence level.

Table 3.

Cross-Section ML Estimates For 1990

Dependent Variable:
Ln AFDC Benefits 1990
for a family of size 3.

Explanatory Variables:	Maximum Likelihood Coefficient Estimates[1]		
	Weight Matrices		
	W_I	W_{II}	W_{III}
W AFDC	0.30 (2.38)	0.22* (1.82)	0.29 (1.13)
Constant	-12.1 (-4.17)	-12.3 (-4.09)	-13.9 (-4.60)
Ln Per Capita Income	1.20* (1.85)	1.33 (2.03)	1.51 (2.24)
Ln State Share Interm. Inc. States [2]	0.33 (0.99)	0.30 (0.88)	0.28 (0.79)
Dummy State Share High Inc. States	12.7 (0.97)	11.6 (0.87)	10.8 (0.78)
Ln Prop. Afro-American	-0.09 (-2.52)	-0.10 (-2.63)	-0.11 (-3.17)
Ln Prop. Dem. State Representative	0.30 (2.59)	0.32 (2.72)	0.32 (2.70)
Ln Female Unemployment	-0.37 (-2.42)	-0.39 (-2.53)	-0.44 (-2.76)
Ln Pct. Recip. Unmarried Parent	-0.32* (-1.82)	-0.35 (-1.99)	-0.37 (-2.06)
Tests of Error Specification[3]			
LM Breusch-Pagan Heterosked.	7.05 (0.42)	7.17 (0.41)	6.41 (0.49)
Kiefer-Salmon Normality Test	0.78 (0.68)	0.66 (0.72)	0.46 (0.79)
LM Test of Spat. Error Depen.	1.93 (0.17)	4.30 (0.04)	2.20 (0.14)
Robust Tests of Spat. Dep.			
LM Test of Spat. Lag Depen.	7.52 (0.00)	8.18 (0.00)	4.0 (0.05)
LM Test of Spat. Error Depen.	3.40 (0.07)	6.47 (0.01)	4.32 (0.04)
Common Factor Test			
LR Test on Common Factor Hyp.	16.2 (0.02)	13.8 (0.05)	8.4 (0.30)

[1] Asymptotic standard normal statistics are given in parenthesis.
[2] The corresponding estimates of the per capita income coefficient for states with matching rate inside the boundaries are: 1.85 (6.15), 1.92 (6.24), 2.06 (6.72), respectively.
[3] *p-values* are given in parenthesis.
* Significant at a 90% confidence level.

Table 4.

Cross-Section ML Estimates For 1995

Dependent Variable: Ln AFDC Benefits 1995 for a family of size 3.	Maximum Likelihood Coefficient Estimates[1]		
	Weight Matrices		
Explanatory Variables:	W_I	W_{II}	W_{III}
W AFDC	0.32 (2.57)	0.31 (2.58)	0.30 (0.26)
Constant	-13.58 (-4.30)	-13.45 (-4.23)	-16.15 (-4.93)
Ln Per Capita Income	1.02* (1.65)	1.06* (1.73)	1.43 (2.19)
Ln State Share Interm. Inc. States[2]	0.45 (1.38)	0.44 (1.37)	0.40 (1.17)
Dummy State Share High Inc. States	18.2 (1.38)	17.8 (1.36)	16.2 (1.16)
Ln Prop. Afro-American	-0.10 (-2.58)	-0.10 (-2.50)	-0.13 (-3.30)
Ln Prop. Dem. State Representative	0.13 (1.46)	0.14 (1.62)	0.16* (1.72)
Ln Female Unemployment	-0.14 (-1.02)	-0.15 (-1.08)	-0.20 (-1.33)
Ln Pct. Recip. Unmarried Parent	-0.41* (-1.71)	-0.46 (-1.92)	-0.49 (-1.96)
Tests of Error Specification[3]			
LM Breusch-Pagan Heterosked.	5.31 (0.62)	5.86 (0.57)	4.73 (0.69)
Kiefer-Salmon Normality Test	0.54 (0.76)	0.46 (0.79)	0.27 (0.87)
LM Test of Spat. Error Depen.	1.25 (0.26)	0.71 (0.40)	0.40 (0.52)
Robust Tests of Spat. Dep.			
LM Test of Spat. Lag Depen.	6.71 (0.00)	5.83 (0.02)	2.14 (0.14)
LM Test of Spat. Error Depen.	2.80 (0.09)	2.30 (0.13)	1.76 (0.18)
Common Factor Test			
LR Test on Common Factor Hyp.	22.8 (0.00)	13.6 (0.06)	13.0 (0.07)

[1] Asymptotic standard normal statistics are given in parenthesis.

[2] The corresponding estimates of the per capita income coefficient for states with matching rate inside the boundaries are: 1.93 (6.12), 1.94 (6.16), 2.23 (6.89), respectively.

[3] *p-values* are given in parenthesis.

* Significant at a 90% confidence level.

Table 5.
Pooled Cross-Section ML Estimates
Years 1985, 1990, and 1995

Dependent Variable: Ln AFDC Benefits	Maximum Likelihood Coefficient Estimates[1]					
	Weight Matrices					
Explanatory Variables:	W_I	W_{II}	W_{III}	W_I	W_{II}	W_{III}
	Non-fixed-Effects			Fixed-Effects		
W AFDC	0.57 (8.81)	0.53 (8.28)	0.80 (13.9)	0.42 (5.14)	0.42 (5.42)	0.72 (7.76)
Constant	-7.20 (-3.64)	-6.90 (-3.41)	-10.7 (-5.09)	5.99 (3.62)	5.94 (3.72)	1.12 (0.71)
Ln Per Capita Income	0.58* (1.78)	0.67 (2.00)	0.91 (2.62)	-0.42 (-1.56)	-0.23 (-1.47)	0.05 (0.31)
Ln State Share Interm. Inc. States[2]	0.27 (1.58)	0.24 (1.40)	0.25 (1.35)	-0.09 (-1.19)	-0.10 (-0.64)	-0.07 (-0.84)
Dummy State Share High Inc. States	10.5 (1.59)	9.4 (1.41)	9.7 (1.36)	-3.27 (-1.13)	-3.51 (-1.23)	-2.22 (-0.80)
Ln Prop. Afro-American	-0.06 (-2.58)	-0.06 (-2.47)	-0.10 (-3.97)	-0.29 (-3.25)	-0.30 (-3.39)	-0.25 (-2.90)
Ln Prop. Dem. State Representative	0.13 (1.94)	0.14 (2.01)	0.13* (1.84)	0.04 (1.10)	0.04 (1.07)	0.02 (0.59)
Ln Female Unemployment	0.08 (0.99)	0.07 (0.85)	0.02 (0.25)	0.03 (0.98)	0.03 (0.97)	0.05* (1.62)
Ln Pct. Recip. Unmarried Parent	-0.33 (-2.69)	-0.40 (-3.09)	-0.46 (-3.41)	-0.05 (-0.59)	-0.05 (-0.64)	-0.03 (-0.40)
Tests of Error Specification[3]						
LM Breusch-Pagan Heterosked.	10.6 (0.16)	13.4 (0.06)	9.5 (0.21)	82.4 (0.00)	84.4 (0.00)	80.4 (0.00)
Kiefer-Salmon Normality Test	1.21 (0.55)	0.15 (0.93)	0.05 (0.98)	0.26 (0.88)	0.42 (0.81)	0.07 (0.96)
LM Test of Spat. Error Depen.	1.81 (0.18)	1.75 (0.19)	21.6 (0.00)	10.45 (0.00)	5.64 (0.02)	3.53 (0.06)
Robust Tests of Spat. Dep.						
LM Test of Spat. Lag Depen.	6.38 (0.01)	5.49 (0.02)	0.88 (0.35)	27.8 (0.00)	30.3 (0.00)	31.8 (0.00)
LM Test of Spat. Error Depen.	2.80 (0.09)	1.18 (0.28)	45.5 (0.00)	12.7 (0.00)	15.6 (0.00)	1.59 (0.21)
Common Factor Test						
LR Test on Common Factor Hyp.[4]	15.4 (0.03)	16.0 (0.03)	20.9 (0.00)	NA -	NA -	NA -

[1] Asymptotic standard normal statistics are given in parenthesis.
[2] The estimates of the per capita income coefficient for states with matching rate inside the boundaries are 1.12 (5.77), 1.14 (5.73), 0.91 (2.62), -0.42 (-3.09), -0.42 (-3.13), and -0.07 (-0.56), respectively.
[3] *p-values* are given in parenthesis.
[4] This test is not computed for the Fixed-Effects models because the variables WX lead to perfect multicollinearity.
* Significant at a 90% confidence level.

Table 6.

Pooled Cross-Section GMM Estimates
Years 1985, 1990, and 1995

Dependent Variable: Ln AFDC Benefits	GMM Coefficient Estimates[1]					
	Weight Matrices					
Explanatory Variables:	W_I	W_{II}	W_{III}	W_I	W_{II}	W_{III}
	Non-fixed-Effects			Fixed-Effects		
W AFDC	0.55 (4.23)	0.39 (2.65)	1.06 (4.59)	1.35 (4.28)	0.94 (3.15)	1.22 (7.23)
Constant	-7.31 (-3.69)	-7.05 (-3.36)	-11.8 (-5.47)	-5.76 (-1.33)	-0.66 (-0.16)	-5.80 (-2.05)
Ln Per Capita Income	0.64* (1.88)	0.79 (2.14)	1.01 (2.79)	0.50* (1.78)	0.23 (0.80)	0.53 (2.55)
Ln State Share Interm. Inc. States[2]	0.26 (1.50)	0.24 (1.27)	0.20 (1.11)	-0.12 (-1.41)	-0.12 (-1.51)	-0.08 (-1.16)
Dummy State Share High Inc. States	10.0 (1.52)	9.20 (1.29)	7.8 (1.10)	-4.43 (-1.39)	-4.63 (-1.49)	-3.13 (-1.13)
Ln Prop. Afro-American	-0.07 (-2.13)	-0.08 (-2.27)	-0.07 (-2.57)	-0.27 (-2.86)	-0.26 (-3.20)	-0.22 (-2.62)
Ln Prop. Dem. State Representative	0.14 (1.96)	0.16 (2.20)	0.13* (1.83)	-0.00 (-0.07)	0.02 (0.54)	-0.00 (-0.14)
Ln Female Unemployment	0.07 (0.75)	0.03 (0.35)	-0.09 (-0.96)	-0.05 (-1.07)	-0.00 (-0.14)	0.04 (1.51)
Ln Pct. Recip. Unmarried Parent	-0.34 (-2.69)	-0.44 (-3.40)	-0.47 (-3.67)	0.02 (0.25)	-0.03 (-0.44)	0.03 (0.40)
Lambda[3]	0.11 (40.9)	0.22 (3.78)	0.31 (8.31)	-0.54 (-18.4)	-0.68 (-2.47)	-1.37 (-2.76)
GMM Tests of Spat. Dep.[4]						
LM_{GMM} *Test of Spat. Lag Depen.*	18.9 (0.00)	9.8 (0.00)	27.6 (0.00)	41.5 (0.00)	44.8 (0.00)	96.2 (0.00)

[1] t-values are given in parenthesis.
[2] The estimates of the per capita income coefficient for states with matching rate inside the boundaries are: 0.64* (1.88), 0.79 (2.14), 1.01 (2.79), 0.27 (1.02), -0.01 (-0.04), and 0.36* (1.80), respectively.
[3] Lambda estimates correspond to the Non-Linear Least Squares estimates.
[4] *p-values* are given in parenthesis.
* Significant at a 90% confidence level.

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Endnotes

[1]Some other strategic variables used in the fiscal competition literature are environmental standards (Oates and Schwab [29]), population growth controls (Brueckner [10]), and composition of public goods (Keen and Marchand [18]).

[2]Peterson et al. [32] argue that even if no welfare-induced migration existed, states still set welfare benefits too low to avoid moral hazard problems like work disincentives, births out of wedlock and non-marriage, which have been widely associated with the cash support program Aid to Families with Dependent Children (AFDC). For a review of the incentive effects of the U.S. welfare system, see Moffitt [24].

[3]The reform, based on the Personal Responsibility and Work Opportunity Reconciliation Act of 1996, gave more responsibility to the states in the provision of welfare, limited the financial role played by the federal government, and imposed work requirements and time limits in the duration of benefits. Under the TANF program the federal government provides states with a fixed "block grant" amount that is independent of the level of state expenditures on welfare.

[4]Unlike in Wildasin [41], we exclude a central government to illustrate the implications of welfare competition with no intervention. Theoretically, the inclusion of a system of matching grants can induce increases in state welfare benefits (see footnote 7). In the empirical tests of AFDC competition we estimate the effect of the federal grants on AFDC benefits.

[5]After a AFDC welfare recipient has been working for more than four months he or she faces a dollar loss in benefits for every dollar increase in earned income.

[6]Note that $\partial Z/\partial b_i = 1 + f''(l_i)\partial l_i/\partial b_i$. Therefore, using (3a), $\partial Z/\partial b_i = 1 - (I-1)/I$, which is greater than zero, but smaller than one.

[7]The theory suggests that a decentralized system of welfare provision needs a price correction mechanism to raise welfare payments, which are in general too low as a consequence of welfare competition. It is argued that a matching grant system can decrease the price of an additional dollar of benefits for states and therefore, can induce increases in state welfare benefits. However, welfare reform put in place a block grant system that, according to the predictions, will induce a decline in the AFDC benefit levels because the marginal cost of welfare spending will be borne entirely by the state governments. See Brueckner [9] and the references therein for a recent discussion of the welfare reform and the "race to the bottom."

[8]Equation (8) is obtained after substituting $l_i = (1/2)L + (b_1 - b_2)/4\alpha$ and $f_i''(l_i) = -2\alpha$ in equation (6a).

[9]In this case the optimal choice of the welfare payment of jurisdiction 1 becomes $b_1 = (1/\gamma)(\delta + \gamma L\alpha - L/2n_i - \beta\gamma)$, where $l_1 = L/2$ has been assumed. Similar equations and conclusions are derived for jurisdiction 2.

[10]Because the functional form of the reaction function depends on the specification of the production and utility functions in the game, it is not possible to obtain structure from (7). For instance, an example like the one described in the text but with Cobb-Douglas utility function leads to a highly non-linear, but non-closed-form representation of the reaction function. We can represent (7) with a local linear approximation to a general non-linear form of the reaction function (see Smith [34]).

[11]See also Case et. al. [13] for a model of spillovers and state expenditures interdependence.

[12]OLS estimates are inconsistent because of the correlation between WB and the disturbance term ε. For more details see Anselin [1], pp. 58-59.

[13]The full version of the model (i.e. (11) and (13) together) can be written as $B = (\phi + \lambda)WB - \lambda\phi W^2 B + X\theta - \lambda WX\theta + \nu$. Note that the coefficients ϕ and λ can be identified from the two different variables WB and W^2B only if non-linear constraints among these parameters and the other parameters of the model are strictly enforced.

[14]Kelejian and Robinson [19] suggested a GMM estimator for a spatial model with spatial error dependence given by $\varepsilon = \lambda W v + \nu$. Because we assume that the error process follows (13) instead, we need to use a different consistent estimator of the error parameters than the one described in their paper. The consistent estimator we use follows from Kelejian and Prucha [20]. The details of the derivation are available upon request.

[15]It is important to note that we use the GMM estimator to draw inferences from a model where spatial error correlation is explicitly modeled.

[16]See Brueckner [9] for a description of the results of these studies.

[17]These authors do not take into account spatial error correlation. This may be a potential problem because spatial error correlation can give spurious evidence of welfare competition.

[18]We excluded the state of Nebraska from our samples because the proportion of state

representatives that are Democrats was not defined for this state. Instead, this state has an unicameral body elected without party designation (see the Statistical Abstract of the United States, U.S. Bureau of the Census, several issues).

[19] Monetary variables like income and AFDC benefits are adjusted for changes in the cost of living over time for the pooled cross-section and fixed-effects estimations, using the CPI. The data comes from several editions of Background Material and Data on Programs within the Jurisdiction of the Committee on Ways and Means, U.S. House of Representatives; CPI Detailed Report (Bureau of Labor Statistics (BLS)); Geographic Profile of Employment and Unemployment (BLS); the Social Security Bulletin (Social security Administration); Statistical Abstracts of the U.S. (U.S. Bureau of the Census); Characteristics and Financial Circumstances of AFDC Recipients of the Family Support Administration (U.S. Department of Health and Human Services, Administration for Children and Families, and Office of Family Assistance); and the Survey of Current Business, 1996 (U.S. Department of Commerce).

[20] Incidence of non-marriage in the population of welfare recipients may be regarded as endogenous. However, there is not convincing evidence that support this hypothesis (see Moffitt [24]). In this paper we treat this variable as exogenous.

[21] The federal share is determined by $(1-S)$, where S is the state share, which has a minimum of 0.50 and a maximum of 0.83. The state share is given by $0.45(Y_s^2)/Y_n^2$, where Y is per capita income, and the subscripts s and n refer to the state and nation, respectively. After 1983, all states used this formula in determining the federal share in state AFDC benefits.

[22] In our cross-section sample no state received the maximum matching rate (83%), and therefore, only the lower boundary (50% rate) is relevant for the proposed method.

[23] All the estimates of the combined effect of the matching rate and the per capita income variable for states with matching rates inside the boundaries are presented in a footnote in each table. The variance for these estimates is computed according to the following formula: $var(\hat{\delta}) = (1/(\partial S/\partial Y))^2\ [var(\hat{\alpha}) + var(\hat{\gamma}) - 2covvar(\hat{\alpha}\hat{\gamma})]$. See footnote 21.

[24] Most of the fixed-effects estimates were significantly different from zero, under all specifications of the weight matrix W. The number of non-significant dummies was eleven, under the specification with W_I; fourteen under W_{II}, and twelve under W_{III}.

[25] After controlling for fixed-effects, Smith [35] also reports a negative effect of per capita income on AFDC benefits. He argues that negative estimates may be reflecting the general downward trend in benefit levels over time, as opposed to the upward trend in real per-capita income.

[26] It is known that the introduction of the fixed effects increases the variance of the coefficient estimates, and therefore, efficiency problems arise in this context.

[27] Values of ϕ that are larger than one indicate that there is an explosive pattern of interaction. In our context, this can be interpreted as evidence that there is a non-stable equilibrium. However, this result is not admissible in the ML estimation since stability constraints are imposed in the optimization problem.