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FISCAL DECENTRALIZATION AND LAND POLICIES

Edited by Gregory K. Ingram and Yu-Hung Hong

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CONTENTS

10

Local Government Finances: The Link Between Intergovernmental Transfers and Net Worth

Luiz de Mello

A large literature shows how the sharing of revenue between different levels of government and the design of intergovernmental transfer schemes affect subnational finances (see de Mello 2000 for a review of the literature). Depending on how shared funds are raised (from a common pool of revenue, for instance) and transfer arrangements are designed (unconditional or special purpose, open- or closed-ended, matched or unmatched, discretionary or formula based, etc.), an increase in transfer receipts may lead to a reduction in subnational government net worth. The basic idea is that transfers reduce the marginal cost of provision to be borne by local taxpayers, especially when financed by a common pool of resources mobilized elsewhere in the economy. This cost shifting discourages local revenue mobilization or induces fiscal profligacy, leading to a buildup of debt in the recipient jurisdiction. Causality, however, may also run in the opposite direction: a fall in net worth may trigger an increase in transfers from higher levels of government. Such is the case when grants are of the ex-post gap-filling type, as with outright bailouts of subnational jurisdictions in financial distress by higher levels of government.

I thank Andrew Dean, Ronald C. Fisher, Yu-Hung Hong, Gregory K. Ingram, Peter Jarrett, Val Koromzay, Diego Moccero, and the conference participants at the Lincoln Institute International Conference for comments and discussions. The views expressed in this chapter are my own and do not necessarily reflect those of the Organisation for Economic Co-operation and Development.

Against this background, using a panel of Organisation for Economic Co-operation and Development (OECD) countries from 1980 through 2005, this study tests for (1) the presence of a stable, long-run statistical association between changes in transfer receipts and subnational net worth; and (2) the direction of causality between changes in transfer receipts and net worth. If a stable long-run association is found to exist and changes in transfer receipts temporally cause changes in net worth, the empirical findings would lend credence to the cost-shifting hypothesis. If causality is found to run in the opposite direction, the results would favor the ex-post soft-budget-constraint hypothesis. In particular, panel-based unit roots and cointegration techniques can be used to test for the existence of a stable relationship between transfer receipts and net worth and, should this relationship exist, to estimate the relevant long-term parameter.

This chapter's main contribution is twofold. First, it fills a gap in the empirical literature by testing for temporal causality in the association between intergovernmental transfers and subnational net worth, with an emphasis on local governments. Although there is a large literature on how intergovernmental transfer arrangements affect subnational finances (reviewed below), the analysis of temporal causality between transfer receipts and subnational net worth is a novelty. Second, attention is shifted away from the use of country-specific budgetary data, which is common in the empirical literature, toward cross-country national accounts data. In doing so, this study aims to highlight statistical regularities that go beyond country-specific institutional arrangements, while dealing with the effect that these arrangements can have on subnational public finances by exploiting heterogeneity in the panel. The main advantage of using national accounts data in the empirical analysis is that they allow for greater cross-country comparability of public finance indicators than do budgetary data, which tend to differ considerably across countries on the basis of differences in coverage and reporting standards.

The following main empirical findings are reported herein:

- There is a stable long-term relationship between transfer receipts and local government net worth for the case of current, but not capital, transfers. The estimated parameter shows that an increase in intergovernmental transfer receipts is associated with a modest reduction in the recipient jurisdiction's net worth over the long term. In addition, a fall in net worth is also associated with an almost one-to-one subsequent increase in transfer receipts.
- The direction of causality is sensitive to the technique used to estimate the long-term parameters. One technique suggests that causality runs from transfers to net worth, which lends support to a large literature on the effect of cost shifting on subnational budget outcomes. Causality also appears to run from net worth to transfer receipts, however, suggesting that transfers may be used as a deficit-financing tool, as when subnational governments are bailed out by higher levels of government.

The Literature

Two main strands of literature suggest a link between intergovernmental transfer receipts and the recipient jurisdiction's indebtedness. One focuses on the association between the design of intergovernmental transfer systems and budget outcomes through cost shifting and predicts that reliance on transfer receipts to finance subnational provision leads to a reduction in subnational indebtedness by weakening incentives for fiscal prudence. Of particular interest in this strand of literature is the "flypaper effect," according to which the "transmission mechanism" between incoming transfers and indebtedness is through expenditure pressures. The other related strand of literature focuses on the effect of soft-budget constraints on subnational finances. Accordingly, higher levels of government may use discretionary grants to bail out lower-level jurisdictions in financial distress. Expectations of financial bailouts reduce the opportunity cost of borrowing, which creates incentives for profligacy. The theoretical underpinning of both strands of literature are therefore that intergovernmental transfers place a wedge between the costs and benefits of local provision, which distorts the incentives faced by local policy makers for fiscal rectitude.

THE DEFICIT-BIAS HYPOTHESIS: TRANSFERS CAUSE INDEBTEDNESS

The basic idea of the deficit-bias literature is that intergovernmental transfer receipts create a wedge between the costs of public provision to be borne by taxpayers in the recipient jurisdiction and the benefits they accrue from public provision, especially when it is financed from a "common pool" of revenue mobilized elsewhere in the economy (Hallerberg and von Hagen 1999; von Hagen and Harden 1995). This wedge allows the recipient jurisdiction to internalize the benefits of expenditure among local residents and to shift provision costs to nonresidents. The upshot is that, due to a range of institutional and political-economy factors, dependence on grants and transfers from higher levels of government creates a deficit bias at the subnational level because it encourages recipient jurisdictions to underutilize their own tax bases at the expense of sharable bases or to spend beyond their means. The incentive to delay fiscal adjustment is another consequence of common-pool financing because individual jurisdictions have limited incentives to act alone and strong incentives to free ride, if the burden of fiscal retrenchment can be shared horizontally across jurisdictional borders and vertically across government levels (Alesina and Drazen 1991; Velasco 1999, 2000).

The deficit-bias hypothesis is conventionally tested in a reduced-form regression setup. The subnational budget balance is regressed on a measure of vertical imbalance, such as the ratio of transfer and grant receipts in revenue, as well as appropriate controls for subnational fiscal stance, such as demographics, terms-of-trade effects, and local income. Despite some variation in the estimating equation, there is plenty of empirical evidence in support of the deficit-bias hypothesis. Cross-country evidence for OECD and non-OECD countries is available from de Mello (1999, 2000) and Rodden (2002), among others. Country-specific evidence is also available: Jones, Sanguinetti, and Tommasi (2000) report evidence of "common pool" incentives for fiscal mismanagement among Argentinean provinces arising not only from intergovernmental revenue-sharing arrangements but also from the political system. Evidence of an association between vertical imbalances and subnational borrowing costs—due to a rising risk premium associated with a subnational deficit bias—is reported by Poterba and Rueben (1997) for U.S. states and de Mello (2001) for OECD and non-OECD countries.

A Special Case: The Flypaper Effect Of particular interest when examining the "transmission mechanisms" through which revenue sharing affects budget outcomes is the flypaper-effect literature, surveyed by Hines and Thaler (1995), among others. This strand of literature is motivated by the observation that an increase in grants and transfer receipts from higher-level jurisdictions often leads to a rise in subnational spending that is higher than that associated with an equivalent hike in local income. This finding is puzzling because the median voter model of taxpayer behavior predicts that, instead, equally sized changes in unconditional grants and in local income should have an equivalent effect on subnational spending. In other words, although theory predicts that changes in transfer receipts or local income would create an identical income effect that would put upward pressure on local spending, this prediction is not always validated by empirical observation.

The flypaper hypothesis is conventionally tested by running reduced-form regressions of subnational spending on receipts of grants and transfers from higher levels of government, local income, and appropriate controls for other determinants of subnational expenditure, such as demographics. The empirical findings available to date suggest that the flypaper effect is stronger for capital than current transfer receipts (Wyckoff 1988), for matching than unconditional transfers (Gamkhar and Oates 1996), and for government spending on "luxury" goods (i.e., culture and urban amenities) than on normal goods (Deller and Maher 2005a). Another important finding is that the flypaper effect is asymmetric in the sense that spending tends to be very responsive to increases in transfer receipts, especially when the level of future transfers is uncertain, and comparatively insensitive to reductions. This finding is confirmed by the empirical evidence reported by Gramlich (1987) for U.S. states, Benton (1992) for U.S. state and local governments, Melo (1996) for Colombian subnational jurisdictions, Heyndels (2001) for Flemish municipalities, and Deller and Maher (2005b) for Wisconsin local governments, among others.[1]

1. There are a number of exceptions. For example, Gamkhar and Oates (1996) use U.S. state and local government data for the period 1953–1991 and show that subnational units respond symmetrically to changes in federal grants, regardless of the type of grant (matching or unconditional). Stine (1994) finds a super-flypaper effect using data for Pennsylvania counties during 1978–1988 in that a reduction in transfers induces the recipient jurisdiction to cut back not only spending but also locally raised revenue.

Although the presence of a flypaper effect is now broadly accepted as a statistical "anomaly" in the public finance literature, empirical evidence has been challenged on several grounds. In particular, the flypaper effect is purported to be due to failure to appropriately deal with the endogeneity of transfer receipts (Knight 2000). The argument is that the level of grants and transfers is affected by the political power of recipient jurisdictions, which, in turn, depends on expenditure pressures at the subnational level. The result is a reverse causality bias in the relationship between transfers and spending; therefore, when transfers are instrumented by variables capturing the political power of receiving jurisdictions (i.e., committee representation, proportion of representatives in the majority party, average tenure of representatives), local income and transfer receipts are found to have similar effects on public spending. Another argument that has been used to challenge the empirical evidence is that the flypaper effect is rather sensitive to the functional form of the estimating equation (Becker 1996). Although there is no a priori reason for sensitivity to functional specifications, empirical evidence is typically stronger for log-linear models than for linear estimating equations.

THE SOFT-BUDGET-CONSTRAINT HYPOTHESIS: INDEBTEDNESS CAUSES TRANSFERS

The basic idea about soft-budget constraints and how they affect local public finances is that expectations of a bailout from higher levels of government reduce the opportunity cost of fiscal profligacy. When subnational jurisdictions are free to borrow, they form expectations about how the central government reacts to their financial stance. Higher-level jurisdictions may be willing to assist local governments financially when the public services they provide benefit the rest of society (Wildasin 1997). Because of these externalities, however, the recipient jurisdiction may face the incentive to spend on items generating benefits that can be internalized among residents, rather than on items with stronger interjurisdictional spillovers. Incentives for bailouts may also be stronger in the case of jurisdictions that are "too big to fail."

If the recipient jurisdiction is deficit-prone and has weak incentives to act responsibly, decentralized fiscal management requires incentives for fiscal prudence; otherwise, local fiscal mismanagement may be detrimental to the system as a whole (Qian and Roland 1998). This macrofinancial spillover effect has been at the core of several subnational financial crises (de Mello 1999, 2000; Prud'homme 1995; Tanzi 1995; Ter-Minassian 1999; among others). Hard-budget constraints, especially in the form of fiscal rules, can be self-imposed, introduced by the central government or complemented by market-based scrutiny. In the absence of these safeguards, subnational financial disarray leads to a buildup of debt, which is often financed through bailouts from higher levels of government. Alternatively, Goodspeed (2002) argues that, although soft-budget constraints reduce the opportunity cost of borrowing, they also increase the cost of future taxes needed to pay off at least part of the incremental debt. Where expectations of higher future taxes mitigate the weak opportunity cost of profligacy, borrowing decisions are efficient, as in the case of hard-budget constraints.

A growing empirical literature looks at the association between intergovernmental transfers and indebtedness. While testing for flypaper-type effects, Levaggi and Zanola (2003) show that recipient jurisdictions respond to a decline in grants and transfers through deficit financing, rather than by hiking locally raised revenue or trimming spending, at least as far as the Italian health care system was concerned during the period from 1989 to 1993. Buettner and Wildasin (2006) focus on a sample of U.S. local governments and show that, especially in the case of large cities, fiscal imbalances are financed essentially by offsetting changes in future expenditures and grants. This evidence suggests that intergovernmental transfers act as a fiscal "cushion" for municipalities, which may, in the case of large cities, indicate a softening of budget constraints. Garcia-Mila, Goodspeed, and McGuire (2001) use data for Spanish regions and find evidence in favor of the soft-budget-constraint hypothesis. Martell and Smith (2004) use U.S. state-level data to test empirically the hypothesis that federal grants affect subnational debt issuance, and whether or not there are asymmetries in this relationship when grants are raised or cut back. The empirical findings suggest a correlation between grants and indebtedness: full-faith and credit debt issuance is reported to be positively correlated with both matching and nonmatching grants, whereas the opposite is true for nonguaranteed debt. The authors nevertheless do not distinguish capital and current transfers when assessing the relationship between transfers and debt.

DISTINGUISHING THE COMPETING HYPOTHESES

The difficulty of distinguishing the deficit-bias hypothesis from the soft-budget-constraint hypothesis is that both are observationally equivalent. A statistically significant coefficient in a reduced-form regression of subnational indebtedness on a measure of vertical imbalances and appropriate controls does not allow the econometrician to distinguish between these hypotheses in the absence of temporal causality testing. The deficit-bias literature assumes that the direction of causality runs from transfer receipts to indebtedness, whereas the opposite is true in the soft-budget-constraint literature. Temporal causality testing has nevertheless not been pursued in the empirical literature.

To shed light on this issue, this study first tests for the presence of a stable, long-term association between transfer receipts and recipient jurisdictions' net worth (discussed below) and then proceeds to test for temporal causality. In particular, the competing hypothesis is tested by the deficit-bias hypothesis.

The Deficit-Bias Hypothesis The deficit-bias hypothesis can be tested by regressing subnational net worth on intergovernmental transfer receipts:

$$(1) \qquad D_{it} = \alpha_i^{DB} + \beta^{DB} T_{it} + \nu_{it}^{DB},$$

where D_{it} and T_{it} denote, respectively, net worth and transfer receipts in jurisdiction i at time t, α_i^{DB} are fixed effects, and ν_{it}^{DB} is an error term.

Equation (1) may include other deterministic elements, such as a time trend. The unit root properties of net worth and transfer receipts will be assessed using conventional panel-based procedures, and cointegration testing will be carried out on the basis of the estimated residuals of equation (1). Two procedures will be used to uncover the long-term parameter (β^{DB}). On the basis of temporal causality testing, the deficit-bias hypothesis will not be rejected if the hypothesis that innovations in transfer receipts affect forecasts of net worth cannot be rejected.

The Soft-Budget-Constraint Hypothesis The soft-budget-constraint hypothesis will be tested by regressing intergovernmental transfer receipts on subnational net worth:

$$(2) \qquad T_{it} = \alpha_i^{SB} + \beta^{SB} D_{it} + \nu_{it}^{SB},$$

where T_{it} and D_{it} denote, respectively, transfer receipts and net worth in jurisdiction i at time t, α_i^{SB} are fixed effects, and ν_{it}^{SB} is an error term.

As in the case of equation (1), equation (2) may include other deterministic elements, such as a time trend. Conventional procedures will be used to assess the unit root properties of the data, to test for cointegration between transfer receipts and net worth, and to uncover the long-term parameter. On the basis of temporal causality testing, the soft-budget-constraint hypothesis will not be rejected if the hypothesis that innovations in net worth affect forecasts of transfer receipts cannot be rejected.

Data and Unit Root/Cointegration Tests

Data are available from the summary public finances accounts included in the OECD national accounts database. Information is available on intergovernmental transfers paid and received, net worth, and total revenue and expenditure for four levels of government (central, middle tier, local, and social security funds). The use of net worth is preferred to gross indebtedness because it takes into account the accumulation of financial assets by the recipient jurisdiction.[2] For example, investment programs financed by the issuance of government debt would leave net worth unchanged because an increase in indebtedness would be matched by an accumulation of assets. That is not the case of an increase in current spending commitments financed through higher indebtedness. Information is not available on the composition of financial liabilities by debt instrument (e.g., general-purpose

2. Net worth is the difference between a jurisdiction's gross financial liabilities, which include debt and other short- and long-term liabilities defined by ESA95/SNA93, and its financial assets, which include cash, bank deposits, loans to the private sector, participation in private-sector companies, holdings in public corporations, and foreign exchange reserves.

or revenue-backed issuances) or on the composition of transfers by type of instrument (e.g., matching or unconditional grants, mandated revenue sharing, discretionary or formula-based transfers). Transfer receipt data can nevertheless be decomposed between current and capital transfers.

For most countries, the public finances time series are relatively short. The central government series are typically longer that those for subnational jurisdictions. At the subnational level, data are more readily available for local governments than for middle-tier jurisdictions. Sample selection was therefore guided primarily by data availability. The largest panel that could be obtained from the database includes 13 countries (or less than one-half of the OECD membership) over the period from 1995 to 2004. The main advantage of using the national accounts database in the empirical analysis is that it allows for greater cross-country comparability of public finances indicators than do budgetary data, which tend to differ considerably across countries because of differences in coverage and reporting standards.

Based on the theoretical argument developed above, the variables of interest are the shares in revenue of transfers received by local governments and their level of indebtedness, measured by the ratio of local government net worth to gross domestic product (GDP). The main descriptive statistics of the variables of interest are reported in table 10.1. For example, current transfers account for 33 percent of local government revenue on average, whereas indebtedness is low, given that net worth is nearly balanced on average. There is considerable variation (as gauged by the standard deviation) in the data in the level of indebtedness and in the share of transfers in revenue, however.

Trends in transfer receipts and local government net worth for all countries in the sample are depicted in figure 10.1. Local government net worth as a proportion of GDP trended upward over the reference period in a number of countries, including Austria, Canada, France, The Netherlands, Spain, and Sweden,

Table 10.1
Descriptive Statistics

	Mean	Standard Deviation	Median	Max.	Min.	Number of Observations
Net worth-to-GDP ratio	–0.01	0.03	–0.02	0.11	–0.07	150
Transfers-to-revenue ratio:						
Total transfers	0.38	0.16	0.37	0.75	0.09	112
Current transfers	0.33	0.14	0.34	0.65	0.05	122
Capital transfers	0.05	0.06	0.04	0.26	0.00	112

Note: The sample spans the period 1995–2004.
Sources: OECD national accounts database and the author's calculations.

Figure 10.1
Indebtedness and Transfers: Local Governments

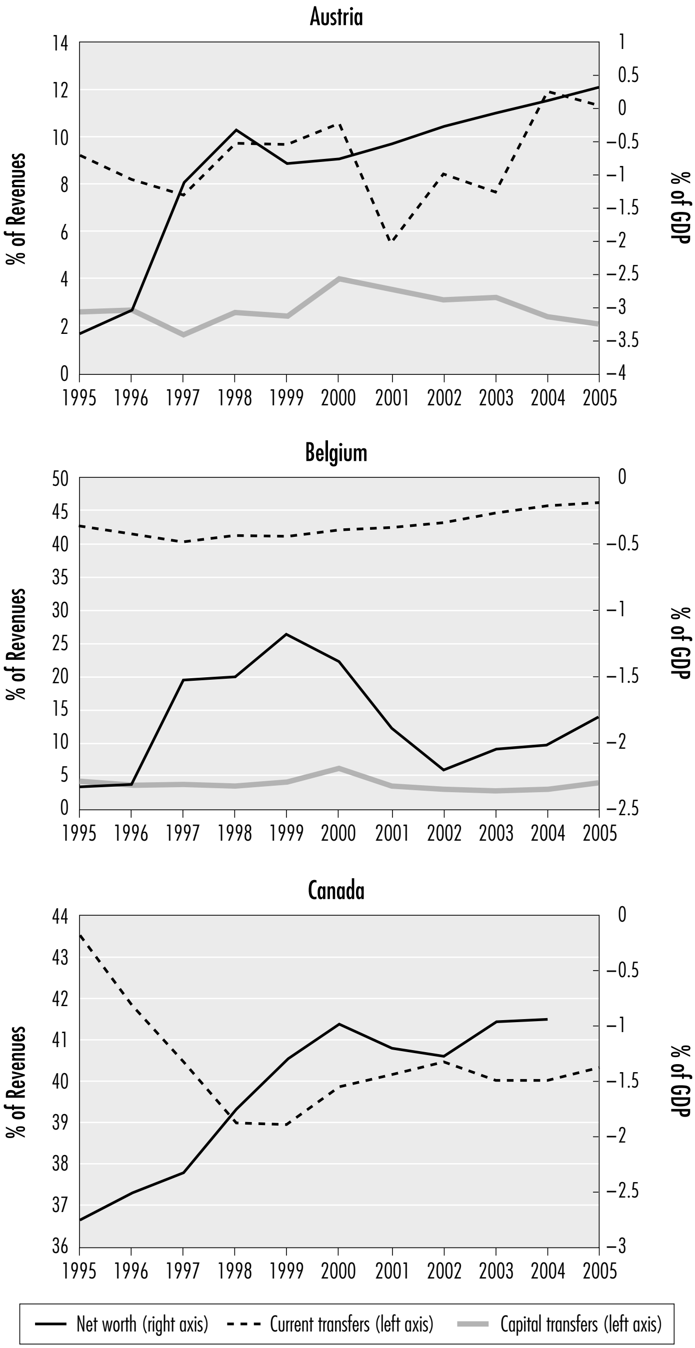


Figure 10.1
(*continued*)

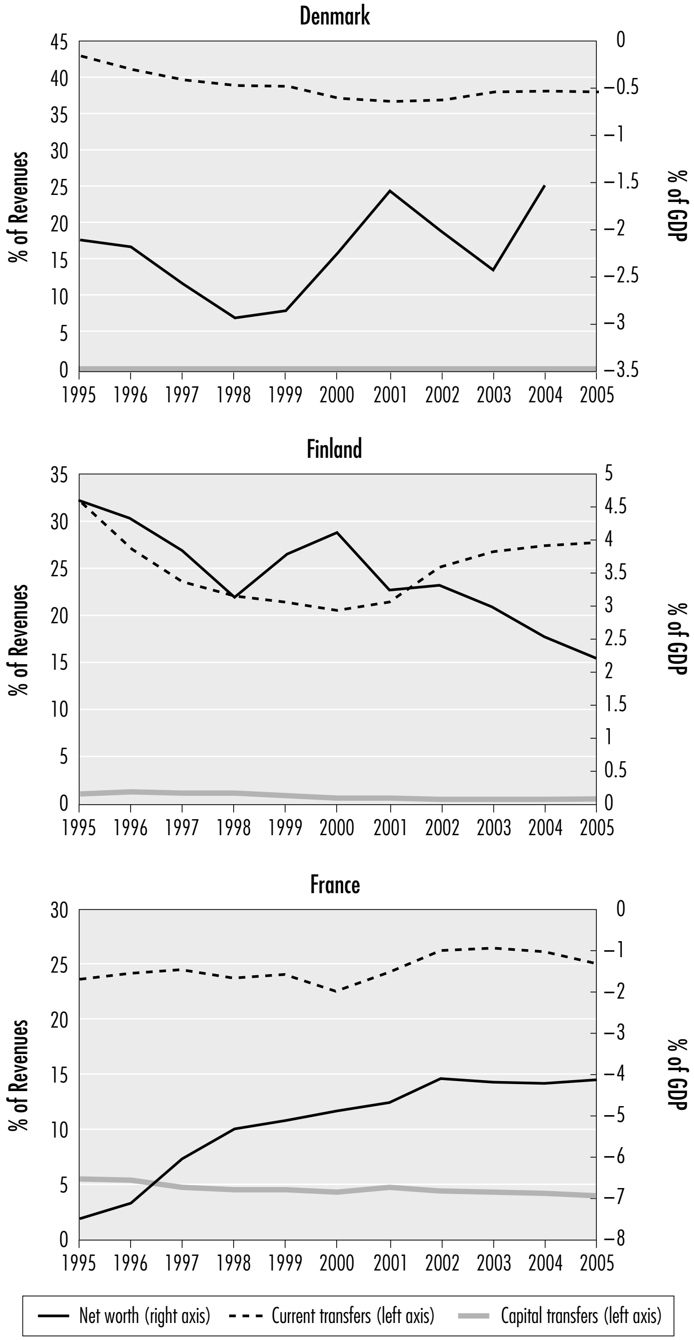


Figure 10.1
(*continued*)

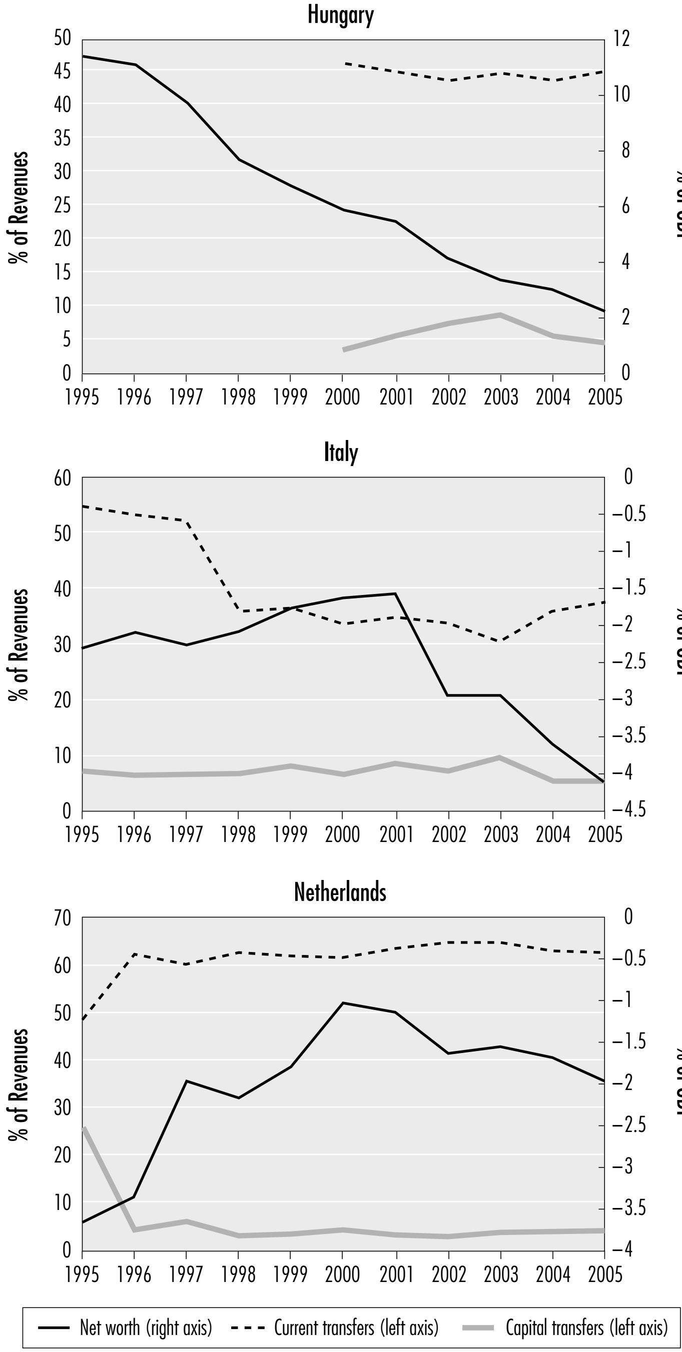
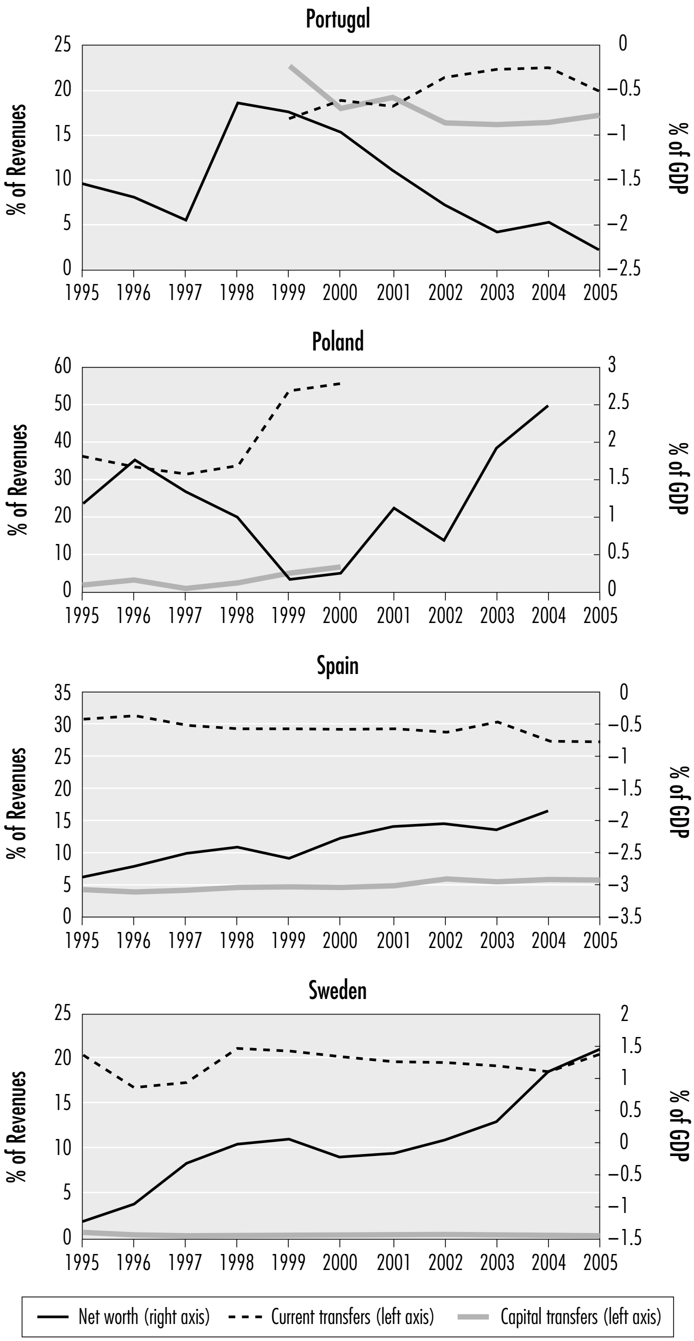


Figure 10.1
(*continued*)



but fell in Finland, Italy, and Portugal. On the other hand, current transfers fell in relation to revenue, albeit often in a gradual manner, in a number of countries, such as Canada, Denmark, Italy, and Spain, while displaying more complex patterns in the remaining countries. The level of capital transfers is typically much lower in relation to revenue in most countries, with the exception of Portugal, and considerably more stable than that of current transfers.

UNIT ROOT TESTS

The unit root properties of the transfer and net worth indicators will be assessed on the basis of four different tests. Three tests are considered for the case in which the cross-sectional units in the panel are independent—Im–Pesaran–Shin (IPS), Maddala–Wu (MW), and Hadri—and one that allows for cross-sectional dependence (CADF).[3] Cross-sectional dependence implies that the time series in the panel are contemporaneously correlated, a phenomenon that may be due to omitted common factors, spatial spillovers, or both. In the case of the variables of interest, the level of subnational indebtedness may be correlated across countries during periods of fiscal retrenchment. An example of such common factor/spatial spillover is the fiscal adjustment effort of the euro-zone countries prior to the common currency's introduction. By the same token, it is important to allow for heterogeneity in the panel when testing for unit roots so that parameter estimates may differ among the different cross-sectional units because the relationship between transfers and indebtedness depends on country-specific institutional settings.

The IPS test is a balanced panel-equivalent of the ADF test with the null hypothesis of a unit root in all cross-sectional units. The alternative hypothesis allows for cross-sectional heterogeneity (i.e., some of the series in the panel are stationary). In other words, rejection of the null hypothesis implies that the variable of interest follows an autoregressive process that contains unit roots in some of the cross-sectional units. The test statistic is a mean-group Lagrange multiplier statistic (t-bar statistic), which converges to a standard normal distribution in large samples (as long as the ratio of N to T tends to a finite nonzero constant as N and T tend to infinity in the case of autocorrelated residuals).

The MW test, proposed by Maddala and Wu (1999), is based on the p-values of individual unit root tests. The null hypothesis is that all series have unit roots, against the alternative that at least one series in the panel is stationary. The MW test differs from the IPS test in the definition of the null hypothesis: it is not based

3. There are several methodologies for testing for unit roots in panel data. Typically, they consist of computing panel-analogs of the Dickey–Fuller (DF) or augmented DF tests available for pure time series, but differ on the definition of the null hypotheses (stationarity or nonstationarity), on whether or not the panel is balanced, and on whether or not heterogeneity is permitted among the autoregressive parameters and across the cross-sections (which affects the definition of the alternative hypotheses). Tests also differ as to whether or not the relevant variables are allowed to be correlated contemporaneously across the cross-sectional units. For recent surveys, see Baltagi and Kao (2000) and Breitung and Pesaran (2006), among others.

on the assumption that the autoregressive coefficient is the same across countries, thus allowing for cross-sectional heterogeneity under the null. The test performs similarly or slightly better than the IPS statistic.

The Hadri test is a panel-equivalent of the KPSS Lagrange multiplier test with the null of stationarity (rather than nonstationarity as in the IPS and MW tests) for all individual series. The error terms may be homoscedastic or heteroscedastic across cross-sectional units, and they may be serially correlated, in which case a Newey–West estimator may be used to take account of the long-run variance in the data. The test nevertheless requires independence across the panel's cross-sectional units and performs poorly in small samples when applied to processes with MA(1) errors.

Finally, the cross-section augmented DF (CADF) test proposed by Pesaran (2005) deals with the case in which cross-sectional dependence arises from the presence of a single common factor among the cross-sectional units. The test averages the individual CADF *t*-statistics for all cross-sectional units in a heterogeneous panel. This test has better size properties than alternative methodologies, such as that proposed by Moon and Perron (2004).

The Results The results of the unit root tests are reported in tables 10.2 and 10.3. The results of the IPS, MW, and Hadri tests suggest the presence of unit roots in net worth (in levels), regardless of whether the disturbances are homoscedastic or not. The results are robust to the inclusion of a time trend in the regressions, where appropriate, given that net worth appears to have a trend in some countries. As for the intergovernmental transfer indicators, the unit root tests yield mixed results. Whereas the transfer-to-revenue ratio appears to have unit roots in levels on the basis of the IPS (except for capital transfers) and the Hadri tests, regardless of whether the disturbances are homoscedastic or not, the MW test suggests that the transfer variables are stationary in levels (except for capital transfers).

On the basis of the results of the CADF test, which allows for contemporaneous correlation among the series in the panel, both the indebtedness and transfer indicators were found to have unit roots in levels. Again, this finding is important because of the comparatively large number of European Union countries in the sample, which creates considerable scope for spatial spillovers and the presence of common factors affecting trends in public finance indicators during the period of analysis.

In sum, the results are unequivocal as to the presence of unit roots in local government net worth, suggesting that the first-differenced data are stationary. Due to the mixed findings for the transfer-to-revenue ratios and the predominance of evidence pointing to the presence of unit roots in the level variables, the cointegration tests (reported below) will be performed on the premise that the transfer indicators are stationary in first differences. Needless to say, a caveat to consider when interpreting the results of the unit root tests is that the time span for which information is currently available is relatively short. It is well known that unit root tests have stronger predictive power when data are available for

Table 10.2
Panel Unit Root Tests: Cross-Sectional Independence

	Level		First Difference	
	Test Statistics	Number of Observations	Test Statistics	Number of Observations
Im–Pesaran–Shin test [H_0: unit root; *t*-bar-statistic]				
Net worth-to-GDP ratio	–1.795	120	–2.341***	105
Transfers-to-revenue ratio				
Total transfers[a]	–1.584	120	–2.477***	90
Current transfers	–1.390	136	–1.996**	119
Capital transfers	–3.981***	120	–2.033**	105
Net transfers[a]	–1.411	112	–2.433***	84
Maddala–Wu test [H_0: unit root; Prob > chi sq]				
Debt-to-GDP ratio	0.520	—	0.000***	—
Transfers-to-revenue ratio				
Total transfers[a]	0.000***	—	0.000***	—
Current transfers	0.000***	—	0.000***	—
Capital transfers	0.106	—	0.000***	—
Net transfers[a]	0.000***	—	0.000***	—
Hadri LM test [H_0: no unit root; *Z*(tau)-statistic]				
Net worth-to-GDP ratio				
Homo	7.188***	150	–1.141	135
Hetero	5.886***	150	–0.107	135
Serial correlation	6.231***	150	6.751***	135
Transfers-to-revenue ratio				
Total transfers[a]				
Homo	6.971***	150	–0.432	120
Hetero	5.463***	150	–0.772	120
Serial correlation	5.413***	150	9.197***	120
Current transfers				
Homo	7.056***	170	–1.531	153
Hetero	4.287***	170	–0.029	153
Serial correlation	6.393***	170	7.025***	153
Capital transfers				
Homo	3.436***	150	0.019	135
Hetero	3.139***	150	0.087	135
Serial correlation	6.051***	150	6.817***	135

(continued)

Table 10.2
(*continued*)

	Level		First Difference	
	Test Statistics	**Number of Observations**	**Test Statistics**	**Number of Observations**
Net transfers[a]				
Homo	7.686***	140	–1.261	126
Hetero	5.338***	140	–0.02	126
Serial correlation	5.451***	140	5.359***	126

* = $p < .10$
** = $p < .05$
*** = $p < .01$
Note: The sample spans the period 1995–2004. The regressions for the IPS test include a constant term, and the variables are lagged once. For the Hadri test, "Homo" and "Hetero" refer, respectively, to the statistics under the hypotheses of homoscedastic and heteroscedastic disturbances across cross-sectional units. The statistics under "Serial correlation" were computed by controlling for autocorrelation in the error terms (lag length is truncated at 2).
[a]Twice-differenced.
Sources: OECD national accounts database and the author's estimations.

Table 10.3
Panel Unit Root Tests: Cross-Sectional Dependence

	Level		First Difference	
	Test Statistics	**Number of Observations**	**Test Statistics**	**Number of Observations**
	CADF (H_0: unit root; *t*-bar-statistic)			
Net worth-to-GDP ratio[a]	–2.088	120	2.610***	90
Transfers-to-revenue ratio				
Total transfers[a]	–1.342	120	2.610***	90
Current transfers	–1.627	136	–1.907	119
Capital transfers	–1.883	120	–2.434**	105
Net transfers[a]	–1.423	112	2.610***	84

* = $p < .10$
** = $p < .05$
*** = $p < .01$
Note: The sample spans the period 1995–2004. The regressions include a constant term.
[a]Twice-differenced.
Sources: OECD national accounts database and the author's estimations.

much longer time periods and when the time dimension of the panel is higher than its cross-sectional dimension.

COINTEGRATION TESTS

A number of methodologies are now available for testing for panel cointegration.[4] As with unit root tests, these methodologies are panel counterparts of pure time-series techniques. One method uses residuals-based approaches akin to those of Engle-Granger, such as the Pedroni (1997, 1999) framework, which allows for unbalanced panels and heterogeneity in the slope coefficients as well as fixed effects and trends in the data. The idea of residuals-based tests is that, as in the pure time-series case, if the estimated residuals are stationary, there exists a linear combination among the variables included in the regression.[5]

Again, as in the case of the unit root tests reported above, it is important to allow for cross-sectional heterogeneity to account for the different institutional and country-specific settings that may affect the relationship between intergovernmental transfer receipts and local government net worth. It is not be possible, however, to deal with the presence of common factors and spatial spillovers when testing for cointegration, as was the case of the unit root test analysis reported above. Although recent developments in panel cointegration testing have focused on techniques that allow for cross-sectional dependence arising from common factors, these methodologies require a much larger time dimension than that of the panel considered here.[6]

The Pedroni methodology consists of testing for the presence of unit roots in the residuals of the cointegrating equation. Seven panel statistics are available: four statistics based on the panel's within dimension (panel-ADF statistics) and three based on the panel's between dimension (group-ADF statistics). The null hypothesis is of no cointegration (i.e., unit roots in the residuals) in all cases (Pedroni, 1999, 2001, 2004). The difference between the panel-ADF and group-ADF statistics is related to the specification of the alternative hypothesis: $H_A : \rho_i = \rho < 1$, for all i, for the panel-ADF statistics (where ρ_i is the autoregressive coefficient in a standard ADF equation for the residuals of the cointegrating equation),

4. See Baltagi and Kao (2000) and Breitung and Pesaran (2006) for recent surveys.

5. When more than one within-group cointregrating relationship may exist, there are rank-based tests akin to that of Johansen-Juselius for pure time series because, as in the pure time-series case, residuals-based tests do not allow for identifying the number of cointegrating relationships that may exist among the integrated variables of interest. Among these tests is the maximum likelihood test of cointegrating rank in heterogeneous panels proposed by Larsson, Lyhagen, and Lothgren (2001). Of course, that is not the case at hand because there can be at most one cointegrating relationship between two variables.

6. The asymptotic equivalence between estimators based on cross-independence and those based on cross-dependence in nonstationary panel time series has been showed by Groen and Kleibergen (1999), who propose a likelihood-based framework for cointegration in panels with a fixed number of error-correction models.

and $H_A : \rho_i < 1$, for all i, for the group-ADF statistics, so that heterogeneity is allowed under the alternative hypothesis. Although the predictive power of these statistics rises with the panel's time-series dimension, the group-ADF and panel-ADF statistics generally perform well in small samples.

In what follows, cointegration will be tested on the basis of one of Pedroni's residuals-based group-ADF statistics. The test allows for heterogeneity in the panel, which is important, as argued above, on the basis of cross-country differences in institutional settings. It involves the calculation of a t-bar statistic (similar to that computed for the IPS unit root test) on the basis of the autoregressive coefficients of standard ADF equations for the residuals of the cointegrating equations estimated for each cross-sectional unit in the panel. The group-ADF statistic is defined as

$$\Psi_{\bar{t}} = \frac{\sqrt{N}\,(\bar{t}_{N,T} - E[\bar{t}_{N,T}\,(p, 0)])}{\sqrt{\mathrm{Var}(\bar{t}_{N,T})}} \Rightarrow N(0, 1),$$

where $\bar{t}_{N,T} = (\sum_{i=1}^{N} t_i)/N$, t_i is the t statistic of each ρ_i in standard ADF equations estimated for the residuals of the cointegration equations estimated for all cross-sectional units in the panel, p is the ADF equation's augmentation order, and $E[\bar{t}_{N,T}(p, 0)]$ and $\mathrm{Var}(\bar{t}_{N,T})$ are the mean and variance of $\bar{t}_{N,T}(p, 0)$ under the null hypothesis of no cointegration ($H_0 : \rho_i = 0$), which were tabulated by Pedroni (1999). The group-ADF statistic diverges to minus infinity under the alternative hypothesis. Therefore, the left tail of the normal distribution is used to assess the critical value for rejecting the null: large negative values imply that the null of no cointegration is rejected.

The Results Because of the need to distinguish between two competing hypotheses (discussed above) and because residuals-based cointegration testing is sensitive to the definition of the cointegrating equation, the group-ADF statistic will be computed for an equation in which net worth is a function of transfer receipts and for another equation in which, conversely, transfer receipts are a function of net worth. The results of the cointegration tests, reported in table 10.4, are not sensitive to the definition of the cointegrating equation. The null of no cointegration was rejected in the case of the current transfer-to-revenue ratio, regardless of the theoretical hypothesis being tested. For the equations including the other transfer indicators, there does not appear to be a common stochastic trend between transfer receipts and local government net worth. On the basis of this test, the long-term coefficients will be estimated for the cointegrating equations defined for local government net worth and current transfer receipts.

ESTIMATING THE COINTEGRATING VECTORS

Having established that a cointegrating relationship exists between the variables of interest, at least for the case of current transfers, the cointegrating vector needs to be estimated under both competing theoretical hypotheses: soft-budget-

Table 10.4
Panel Cointegration Tests: Group-ADF Statistics

Transfer Type	Based on Residuals From:	
	Equation (1)	Equation (2)
Total transfers	1.58	1.00
Current transfers	–3.61***	–2.79***
Capital transfers	4.73	–0.51
Net transfers	2.33	0.93

* = $p < .10$
** = $p < .05$
*** = $p < .01$
Note: The sample spans the period 1995–2004. The regressions include an intercept and a time trend.
Sources: OECD national accounts database and the author's estimations.

constraint and deficit bias. Two methodologies will be used in either case: the dynamic OLS (DOLS) and the dynamic fixed-effects (DFE) estimators. Both techniques assume that the cointegrating vectors are identical for all panel units, and neither allows for cross-sectional dependence. The dynamic seemingly unrelated regressions (DSUR) estimator of Mark, Ogaki, and Sul (2005) and Moon and Perron (2004) allow for cross-sectional dependence in the estimation of the cointegration vector. Unlike DOLS, the DSUR estimator exploits the presence of long-run cross-sectional correlation in the equilibrium errors, which makes it more efficient (Westerlund 2005). DSUR, however, is only feasible for panels in which the number of cross-sectional units is significantly smaller than the time-series dimension, and that is not the case for the panel at hand.

The DOLS Estimator The DOLS estimator, developed by Saikkonen (1991) and Stock and Watson (1993), uses leads and lags of the differenced right-hand-side variable to correct for possible serial correlation and weak exogeneity in a cointegrated regression.[7] Based on equation (1), under the deficit-bias hypothesis, the DOLS equation is defined as

$$D_{it} = \alpha_i^{DB} + \beta_{DOLS}^{DB} T_{i,t-1} + \sum_{j=1}^{p_1} \xi_j^{DB}\, \Delta T_{i,t-j} + \sum_{j=1}^{p_2} \psi_j^{DB}\, \Delta T_{i,t+j} + u_{it}^{DB}. \tag{3}$$

Likewise, based on equation (2), in the case of the soft-budget-constraint hypothesis, the DOLS equation is as follows:

7. The DOLS technique, as well as the fully modified estimator of Phillips and Hansen (1990), produces estimators that are asymptotically normally distributed with zero means (Kao and Chiang 1999).

(4) $$T_{it} = \alpha_i^{SB} + \beta_{DOLS}^{SB} D_{i,t-1} + \sum_{j=1}^{p_1} \xi_j^{SB} \Delta D_{i,t-j} + \sum_{j=1}^{p_2} \psi_j^{SB} \Delta D_{i,t+j} + u_{it}^{SB}.$$

The DFE Estimator The DFE estimator is based on an autoregressive distributed lag (ADRL) model in the case of pure time series (Pesaran and Shin 1999). Under the deficit-bias hypothesis, the DFE methodology involves the estimation of the following model:

(5) $$\Delta D_{it} = \sigma_i^{DB} d_i' + \lambda^{DB} D_{i,t-1} + \beta_{DFE}^{DB} T_{i,t-1} + \sum_{j=1}^{p_1} \omega_j^{DB} \Delta D_{i,t-j} + \sum_{j=1}^{p_2} \phi_j^{DB} \Delta T_{i,t-j} + u_{it}^{DB},$$

where d_i' is a vector of time-invariant regressors.

Likewise, in the case of the soft-budget-constraint hypothesis, the DFE equation is as follows:

(6) $$\Delta T_{it} = \sigma_i^{SB} d_i' + \lambda^{SB} T_{i,t-1} + \beta_{DFE}^{SB} D_{i,t-1} + \sum_{j=1}^{p_1} \omega_j^{SB} \Delta T_{i,t-j} + \sum_{j=1}^{p_2} \phi_j^{SB} \Delta D_{i,t-j} + u_{it}^{SB}.$$

The estimate of the long-run coefficients are given by $\theta_n^{DFE} = -\hat{\beta}_{DFE}^n / \hat{\lambda}^n$, where $\hat{\beta}_{DFE}^n$ and $\hat{\lambda}^n$ are the DFE estimators of β_{DFE}^n and λ^n, for $n = (DB, SB)$ in equations (5) and (6). As mentioned above, the long-term parameter is identical for all cross-sectional units.

Testing for Temporal Causality Both methodologies used to estimate the cointegrating vector lend themselves to temporal causality analysis. As argued above, temporal causality allows for distinguishing the deficit-bias and soft-budget-constraint hypotheses about an association between intergovernmental transfer arrangements and recipient jurisdiction indebtedness. It can be tested using a conventional *F*-test. For example, by equation (5), if $H_0 : \beta_{DFE}^{DB} = \phi_j^{DB} = 0$ is rejected for all j, then T_{it} Granger causes D_{it}, which is in support of the deficit-bias hypothesis. Likewise, by equation (6), if $H_0 : \beta_{DFE}^{SB} = \phi_j^{SB} = 0$ is rejected for all j, then D_{it} Granger causes T_{it}, which is in support of the soft-budget-constraint hypothesis.

Temporal causality can be tested in the alternative setting proposed by Hurlin and Venet (2001) for panels with fixed coefficients. For instance, in the case of the deficit-bias hypothesis, temporal causality testing involves the estimation of the following equation:

(7) $$\Delta D_{it} = \alpha_i^{DB} + \sum_{j=1}^{p_1} \lambda_{(j)}^{DB} \Delta D_{i,t-j} + \sum_{j=0}^{p_2} \xi_{i(j)}^{DB} \Delta T_{i,t-j} + u_{it}^{DB}.$$

Two hypotheses can be considered for the case with homogeneous autoregressive processes:[8] homogeneous noncausality (HNC) and homogeneous causality (HC). The null hypothesis under HNC is $H_0 : \xi^{DB}_{i(j)} = 0$, for all i and j, which is tested against $H_A : \xi^{DB}_{i(j)} \neq 0$, for at least some i and j. Acceptance of the null therefore indicates that transfers do not Granger cause net worth for all cross-sectional units in the panel. Rejection of the null hypothesis indicates instead that for at least one or more units, transfers Granger cause net worth. The HNC statistic is computed by comparing the sum of squared residuals of the unrestricted model in equation (7) (RSS_u) with the sum of squared residuals of a restricted model where the slope coefficients and lags of $\xi^{DB}_{i(j)}\, \Delta T_{i,t-j}$ are set to zero, leaving only the fixed effects and the lags of the dependent variable to predict current values of ΔD_{it} (RSS^{HNC}_r). The HNC test statistic is computed as

$$(8) \qquad F_{HNC} = \frac{(RSS^{HNC}_r - RSS_u) / Np}{RSS_u / [NT - N(1 + p) - p]},$$

where N, p, and T are, respectively, the cross-sectional dimension of the panel, the number of lags used in equation (7), and the time-series dimension of the panel.

Acceptance of the null on the basis of an F-test distributed $[Np, NT - N(1 + p) - p]$ calls for testing the hypothesis of homogeneous causality (HC). The null hypothesis for HC is $H_0 : \xi^{DB}_{i(j)} = \xi^{DB}_{j} \neq 0$, for all i and some j, which is tested against $H_A : \xi^{DB}_{i(j)} \neq \xi^{DB}_{i}$, for at least some i and some j. Acceptance of the null indicates that all cross-sectional units follow the same causal process. The HC test statistic is calculated using the sum of squared residuals from the unrestricted model described above (RSS_u) and the sum of squared residuals of a restricted model in which the slope terms are constrained to equality for all cross-sectional units (RSS^{HC}_r). The HC test statistic is computed as

$$(9) \qquad F_{HNC} = \frac{(RSS^{HC}_r - RSS_u) / p(N - 1)}{RSS_u / [NT - N(1 + p) - p]}.$$

In the case of the soft-budget-constraint hypothesis, the Hurlin–Venet setting involves estimating the following equation:

$$(10) \qquad \Delta T_{it} = \alpha^{SB}_i + \sum_{j=1}^{p_1} \lambda^{SB}_{(j)}\, \Delta T_{i,t-j} + \sum_{j=0}^{p_2} \beta^{SB}_{i(j)}\, \Delta D_{i,t-j} + u^{SB}_{it}.$$

The HNC and HC statistics can therefore be computed using equations (8) and (9) to test for temporal causality.

8. The same statistics can be calculated for each cross-sectional unit so as to allow for heterogeneity arising from different autoregressive processes, but that case will not be considered here.

The Estimated Vectors and Temporal Causality Tests The results of the DOLS and DFE estimations are reported in table 10.5. In the case of the deficit-bias hypothesis, the magnitude and sign of the coefficients confirm the hypothesis that an increase in current transfer receipts from higher levels of government is associated with a decrease in the recipient jurisdiction's net worth over the long term. The estimated coefficient is small in size and only significant at classical levels in the DFE regression, however. In the case of the soft-budget-constraint hypothesis, the cointegrating vector implies that a fall in local government net

Table 10.5
Cointegration Vectors

Hypothesis	**Coefficient**	***N***	**Number of Lags**	**R^2 (within)**	***F***
	Deficit-bias hypothesis				
DOLS regression	–0.03 (0.028)	120 [14]	1	0.01	0.58
H_0: Transfers do not cause net worth (Prob > *F*)	0.317				
DFE regression	–0.04*** (0.018)	106 [14]	2	0.15	2.64**
Implied *LR* coefficient	–0.24				
H_0: Transfers do not cause net worth (Prob > *F*)	0.063**				
	Soft-budget-constraint hypothesis				
DOLS regression	–1.06** (0.509)	95 [14]	2	0.13	2.32**
H_0: Net worth does not cause transfers (Prob > *F*)	0.311				
DFE regression	–1.12*** (0.314)	106 [14]	2	0.51	15.03***
Implied *LR* coefficient	–1.60				
H_0: Net worth does not cause transfers (Prob > *F*)	0.003***				

* = $p < .10$
** = $p < .05$
*** = $p < .01$

Note: The sample spans the period 1995-2004. The coefficients reported are, respectively, $\hat{\beta}^n_{DOLS}$ and $\hat{\beta}^n_{DFE}$ for $n = (DB, SB)$, estimated in equations (3) through (6). The *LR* coefficients are computed as $\theta^{DFE}_n = -(\hat{\beta}^n_{DFE} / \hat{\lambda}^n)\ \hat{\lambda}^n$, for $\hat{\lambda}^n$ estimated in equations (5) and (6). All models include an intercept and fixed effects (not reported). Standard errors are reported in parentheses. The number of cross-section units is reported in brackets. The number of lags and leads was selected on the basis of the Akaike Information Criterion (AIC).
Sources: OECD national accounts database and the author's estimations.

Table 10.6
Temporal Causality Tests: Hurlin–Venet Methodology

	Hypotheses	
	Transfers Do Not Cause Net Worth	**Net Worth Does Not Cause Transfers**
F_{HNC}	1.74***	1.12
F_{HC}	1.58***	0.78

* = $p < .10$
** = $p < .05$
*** = $p < .01$
Note: The sample spans the period 1995–2004. The test statistics are described in equations (8) and (9). HNC and HC refer, respectively, to "homogeneous noncausality" and "homogeneous causality." All models include an intercept and fixed effects (not reported). The number of lags and leads was selected on the basis of the Akaike Information Criterion (AIC).
Sources: OECD national accounts database and the author's estimations.

worth is associated with an almost one-to-one increase in current transfer receipts. Evaluated at the sample means, the coefficients estimated by both DOLS and DFE imply that a fall in the ratio of local government net worth from the current level of near balance to about 5 percent of GDP is associated with an increase in transfer receipts from the current level of 33 percent of local government revenue to about 37 percent.

The results of the temporal causality tests, for both the DOLS and DFE equations, are also reported in table 10.5. On the basis of these tests, it appears that transfer receipts do cause net worth in the temporal causality sense in the DFE equation, when the long-term coefficients are estimated by DFE, which supports the deficit-bias hypothesis. Nevertheless, it also appears that net worth causes transfer receipts on the base of the DFE regression, which is in accordance with the soft-budget-constraint hypothesis. The results of the Hurlin–Venet temporal causality tests are reported in table 10.6. On the basis of these tests, there appears to be support for the deficit-bias hypothesis because the null hypothesis that transfers do not Granger cause net worth is rejected comfortably for all cross-sectional units in the panel. There is nevertheless heterogeneity in the panel on the basis of the HC test because the null that all cross-sectional units follow the same causal process is also comfortably rejected.

Summary of the Main Findings and Discussion

This study used OECD national accounts data to shed additional light on the empirical association between intergovernmental transfer arrangements and subnational public finances. In particular, temporal causality analysis was used to distinguish between the deficit-bias and the soft-budget-constraint hypothesis that underscore the empirical association between intergovernmental transfer receipts and recipient jurisdictions' indebtedness (controlling for the accumulation of financial assets). As noted above, the predictions of the deficit-bias and soft-budget-constraint literatures are otherwise observationally equivalent because a statistical association between transfer receipts and net worth is a necessary condition for both predictions. Although the estimation of the long-term parameters by DFE appears to support the deficit-bias hypothesis, there is equally compelling evidence in favor of the soft-budget-constraint hypothesis in the sample of countries under examination. In this latter case, transfer arrangements may act as an alternative financing mechanism for reducing subnational net indebtedness.

The magnitude of the estimated parameters nevertheless suggests that, although an increase in the share of current transfer receipts in local government revenue leads to a modest deterioration in net worth over the long term, a deterioration in local government net worth is associated with a sizable increment in its current transfer receipts (in percent of revenue). To the extent that this finding indicates budget constraints are less hard than possibly desirable, at least as far the OECD countries in the sample are concerned, there is scope for strengthening subnational budget constraints further. One option for doing so is the introduction of fiscal rules, including administrative controls, such as the need for central government approval of subnational borrowing as in Ireland, Japan, Korea, and the United Kingdom.[9] In some countries, including Mexico, local governments are banned from borrowing abroad.

More comprehensive fiscal rules include ceilings on public debt or debt service, expenditure, or budget balances. Golden rules (i.e., budgeted deficits must not exceed investment spending) are in place in some cases (Germany, Switzerland, and the United Kingdom), although other countries (Hungary, Poland, and Portugal) impose ceilings on the public debt or debt service outlays. Outside the OECD area, the experience of Brazil with fiscal rules is instructive because the successful implementation of comprehensive fiscal responsibility legislation has been instrumental in the country's process of fiscal adjustment since the mid-1990s. Also, markets appear to be a poor substitute for fiscal rules, particularly at the subnational level of government, but have complemented fiscal rules in many cases, such as in Canada and the United States. Finally, international experience suggests that, where in place, attention is needed to avoid fiscal gimmickry as a means of bypassing legal restrictions on borrowing. Common mechanisms include channeling expenditures through the tax system, creating off-budgetary funds, and committing government resources through public-private partnerships and loan guarantees, among others.

In addition, a negative association between transfer receipts and net worth may be unrelated to the cost-shifting incentives and their effect on subnational fiscal performance through soft-budget constraints. Such an association may be

9. See OECD (2002, chap. IV; 2003, chap. V) for more information on OECD countries and de Mello (2007) for the case of Brazil.

due instead to different financing mechanisms that are available for subnational governments, such as, for example, securing future revenue from intergovernmental grants. This operation may be an alternative to pay-as-you-go financing of investment projects, for example. In the United States, municipal bonds can be of two types: general obligation (GO), which are backed by general taxation, and revenue bonds, which are financed by receipts of future taxes, fees, lease payments, federal grants, lottery earnings, or tobacco settlement payments. Whereas issuance of GO bonds is often subject to constitutional limits, such is not the case of revenue bonds. An example outside the OECD area is that of the Brazilian states, which resorted to a "revenue anticipation" instrument extensively, including as a deficit-financing tool, until its use was curtailed as a means of reining in subnational indebtedness.[10]

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10. See Afonso and de Mello (2002) for more information.

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