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Public Sector Wages and Bureaucratic Quality: Evidence from Latin America

Public sector employment is commonly perceived as being less demanding than private sector employment. In many countries, the public sector offers a noncompetitive work environment and a level of job security that cannot be matched in the private sector. Quinn, for instance, finds that private sector workers tend to be subject to more workplace disamenities (such as pace of work, supervision, and workplace hazards) than their public sector counterparts.[1] Bellante and Long show that the fringe benefits enjoyed by public sector workers are substantially larger than those offered in the private sector.[2] Poterba and Rueben show that in 1993 benefits averaged 43.8 percent of wages for U.S. public sector workers and 40.3 percent for U.S. private sector workers.[3] At the same time, it is commonly believed that these advantages of public sector employment are offset by lower average wages. This situation is summed up in a joke in which employees of state-owned enterprises in the former Soviet Union describe their employment arrangement as follows: "We pretend to work, and they pretend to pay us." This does not actually seem to

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1. Quinn (1982).
2. Bellante and Long (1981).
3. Poterba and Rueben (1998).

be the case, however. Mounting evidence indicates that in many countries, wages paid in the public sector are not lower and are often significantly higher than wages paid in the private sector.[4] In essence, there seems to be a public sector wage premium.

The purpose of this paper is twofold. First, the paper presents evidence on public-private wage differentials for a sample of seventeen Latin American countries covering 88 percent of the total population of the region. Second, the paper discusses the relationship between the public-private wage differential and the quality of the public sector and presents evidence for a sample of Latin American countries.

The first part of the paper aims at documenting public-private wage differentials for a sample of Latin American countries over the 1980–98 period. While there exists a large literature testing for the presence of a public sector wage premium in the United States (Ehrenberg and Schwarz survey twenty-three studies that support the presence of a public sector wage premium), very little work has been done at the cross-country level.[5] One exception is Blanchflower, who estimates public-private wage differentials for fifteen member countries of the Organization for Economic Cooperation and Development (OECD). He finds that eleven of the countries have a positive and statistically significant public sector wage premium, while only one (Norway) has a statistically significant public sector wage penalty.[6] Gregory and Borland, who survey more than thirty-four studies, find that the public sector wage premium is high for women, but it is often not statistically significant for men.[7] Panizza and Qiang find a significant public sector premium for both men and women in several Latin American countries.[8]

One interesting issue is the correlation between the wage premium and skills. Katz and Krueger run separate regressions for U.S. workers belonging to different educational groups and show that college-educated males face a public sector penalty, while males with lower education enjoy a public sector premium. (Similar findings are obtained for female workers).[9]

4. See, for instance, Gregory and Borland (1999); Panizza and Qiang (1999).
5. Ehrenberg and Schwarz (1986).
6. Blanchflower (1996).
7. In contrast, Bardasi (1998) finds a significant premium for both men and women working in the Italian public sector.
8. Panizza and Qiang (1999).
9. Katz and Krueger (1991).

Quantile regression analysis applied to U.K. data reveals that the public sector premium is inversely related to an employee's position in the distribution of earnings.[10] Similar results are found for Australia and Sweden.[11] The more concentrated earning distribution in the public sector may have an important selection effect. If the sorting of employees between the public and private sectors is due to unobserved productivity-related characteristics, ordinary least squares (OLS) estimations of the public-private wage differential could yield biased results.[12] In van der Gaag and Vijverberg's study of public-private pay differentials in the Ivory Coast, OLS estimates yield a large public sector premium, but this premium disappears when the authors control for the selection bias.[13]

Even with this last caveat, the existing empirical literature on public-private wage differentials can be summarized by the following three stylized facts: a rent is associated with public sector jobs, in that public sector workers often receive both higher wages and higher nonwage compensation; the wage premium is higher for women than it is for men; and the premium is inversely correlated with earnings (and it may become a penalty for high earners). These general results hold for most of the Latin American countries studied in this paper.

While this paper does not try to provide an explanation for the existence of a public sector premium, the theoretical literature on public-private wage differentials emphasizes the role of the higher level of unionization and the soft budget constraint faced by the public sector.[14] Holmlund, for instance, documents the fact that whereas private sector unions may fully internalize the cost of any increase of the wage bill (through a decrease in employment), public sector unions are able to discharge part of the burden on the private sector.[15] In an efficiency wage setting, the public sector premium could also arise from the higher firing costs faced by the public sec-

10. Disney and others (1997).

11. Gregory and Borland (1999).

12. Katz and Krueger (1991) find that applications for federal government jobs have increased for blue-collar workers and decreased for white-collar workers.

13. Van der Gaag and Vijverberg (1988). Similar results are presented in studies of public-private wage differentials in the Netherlands (van Ophem, 1993; Hartog and Oosterbeek, 1993), Peru (Stelcner, van der Gaag, and Vijverberg, 1989), and Haiti (Terrell, 1993).

14. Ehrenberg and Schwarz (1986) and Gregory and Borland (1999) are two excellent surveys.

15. Holmlund (1993).

tor. These high firing costs tighten the no-shirking constraint for the public sector and lead to higher wages. Hence, higher job security, instead of being compensated by lower wages, could be the main cause of the public sector premium.[16]

The second objective of this paper is to study the relation between public-private wage differentials and the efficiency of the public sector. Higher public sector wages are commonly believed to decrease corruption and increase the efficiency of the public sector.[17] However, empirical evidence has not been kind to the idea that public sector wages and the quality of the public sector (often measured as the level of corruption) are positively related. Three recent papers tackle this issue and either do not find any significant correlation between these two variables or find a statistically but not economically significant correlation.[18] Van Rijckeghem and Weder find that quasi-eradication of corruption would require average public sector wages to be two to eight times higher than average manufacturing wages.[19] As Easterly points out, "People respond to incentives. People respond to incentives. People respond to incentives."[20] Higher wages alone, however, do not seem to be the best method to provide public sector workers with the appropriate set of incentives. Di Tella and Schargrodsky hold that higher public sector wages need to be accompanied by better auditing practices, and Rauch and Evans find that meritocratic recruitment is the most important determinant of the quality of the public sector.[21] For Latin America, I find no clear correlation between the average public-private wage differential and the quality of the public sector, whereas the degree of meritocracy in the public sector (relative to the private sector) and the quality of the public sector demonstrate a positive correlation.

The paper is organized as follows. The next section presents evidence for public-private wage differentials in Latin America. The paper then examines the correlation between public-private wage differentials and efficiency of the public sector. The final section concludes.

16. See, for instance, Panizza (1998). Rodrik (1997) discusses an alternative view, in which high levels of public sector employment are not driven by rents, but rather constitute a method of smoothing income and consumption risk in countries with high levels of output volatility.

17. See, for instance, Nunberg and Nellis (1995).

18. Rauch and Evans (2000); Treisman (2000); Van Rijckeghem and Weder (1997).

19. Van Rijckeghem and Weder (1997).

20. Easterly (1998).

21. Di Tella and Schargrodsky (2000); Rauch and Evans (2000).

Public-Private Wage Differentials in Latin America

This section computes wage differentials between the public and private sectors using data from sixty household surveys covering seventeen Latin American countries over the period 1981–98. For each survey, the paper estimates public sector premiums for men and women with different educational levels. The purpose of this section is twofold: to document public-private wage differentials in a large set of Latin American and Caribbean countries and to generate a data set that facilitates the study of the relation between public sector wages and performance.

The Data

Although several developing countries have good household surveys, different methodologies in data collection and different definitions of the variables make cross-country comparisons extremely difficult. This paper uses household surveys that have uniform coding in their questions on human capital investment and labor market participation. For some countries (namely, the Dominican Republic, Guatemala, Mexico, and Nicaragua), I was only able to find one survey in which it was possible to identify workers employed in the public sector. In other cases, I was able to use several surveys covering the 1980s and 1990s. Table 1 reports the countries studied in this paper and the years in which the surveys were collected.

Almost all the surveys feature national coverage (except Uruguay, which is an urban survey), and they code the definition of income and employment in a similar way (see Székely and others for a detailed description of the surveys used in this paper).[22] A potential problem with the data is that not all the household surveys report the hourly income of the primary job. For instance, surveys in Bolivia and Colombia only report the average hourly income of all jobs (the two values are different for individuals who moonlight). Whenever data are available, I use the hourly wage of the primary job. Panizza and Qiang rerun all the regressions using the average hourly income of all jobs; the results indicate no substantial differences between the two definitions of wage.[23] To mitigate selection bias stemming from school attendance, I only use individuals aged 20 to

22. Székely and others (2000). Data for urban samples are available on request.

TABLE 1. Surveys Used in the Analysis

Country	*Year of survey*
Bolivia	1990, 1993, 1996, 1997
Brazil	1992, 1993, 1995, 1996, 1997
Chile	1987, 1996
Colombia	1990, 1991, 1993, 1997, 1998, 1999
Costa Rica	1983, 1985, 1987, 1991, 1993, 1995, 1997
Dominican Republic	1996
Ecuador	1995, 1998
El Salvador	1995, 1997, 1998
Guatemala	1998
Honduras	1989, 1992, 1996, 1997, 1998
Mexico	1994
Nicaragua	1993
Panama	1979, 1991, 1995, 1997
Paraguay	1995, 1998
Peru	1985, 1994, 1996, 1997
Uruguay	1981, 1989, 1992, 1995, 1997
Venezuela	1981, 1983, 1986, 1989, 1993, 1995, 1997

65. To avoid outliers owing to mistakes in data entry, I rank workers according to their wages and dropped the top and bottom 0.5 percent of the sample.

Table 2 describes the composition of employment in the seventeen Latin American countries studied in this paper. While 13 percent of employed people, on average, work for the public sector, approximately 25 percent of men and 35 percent of women with high education are thus employed. (Given that my sample consists of middle income countries, I define high education as having completed secondary school.) In the formal sector, 30 percent of men and 40 percent of women are employed by the public sector, which absorbs more than 40 percent of formal sector workers with high school or university degrees.

Of course, the seventeen countries studied in this paper are far from homogeneous. Chile, Colombia, Ecuador, Guatemala, Honduras, and Paraguay have relatively small public sectors, while Costa Rica, Nicaragua, Panama, and Venezuela have large public sectors that employ more than 20 percent of the workers. Although within-country variation seems to be less important than cross-country variation, many countries

23. Panizza and Qiang (1999).

TABLE 2. Composition of Employment in Latin America[a]

Percent workers employed in the public sector

		Men			Women		
Period	*All workers*	*All men*	*High education*	*Low education*	*All women*	*High education*	*Low education*
All sectors							
All years	14.5	12.7	26.2	8.3	18.1	36.8	8.4
1993–99	13.4	11.6	24.1	7.2	17.0	34.8	7.1
Formal sector only							
All years	31.2	26.6	40.6	19.2	40.3	52.0	25.6
1993–99	29.5	24.8	38.3	17.3	38.8	50.0	23.6

Source: Author's calculations.

a. Unweighted regional averages. Negative values indicate a public sector penalty.

experienced a dramatic reduction in the share of workers employed in the public sector during the sample period. The most extreme cases are Bolivia (where the share of public sector workers dropped from 21 percent in 1990 to 12 percent in 1997), Costa Rica (from 24 percent in 1983 to 16 percent in 1997), Honduras (from 13 percent in 1989 to 9 percent in 1998), Panama (from 27 percent in 1991 to 22 percent in 1997), Uruguay (from 25 percent in 1989 to 19 percent in 1997), and Venezuela (from 23 percent in 1981 to 16 percent in 1997).

Estimations

This paper measures wage differentials between public and private sector workers using a standard dummy variable approach. This approach consists of modeling the effect of sector as an intercept effect, while holding returns to other productivity-related characteristics and job attributes equal across sectors. Formally, I estimate public-private wage differentials using the following specification:

$$w_i = \mathbf{X}_i\boldsymbol{\beta} + P_i\gamma + u_i, \tag{1}$$

where w_i is the log of hourly earnings, $\mathbf{X}_i$ is a matrix of productivity-related characteristics, P_i is a dummy variable that takes a value of 1 when the employee works for the public sector and 0 otherwise, and γ measures the public-private pay differential.[24] I estimate equation 1 separately for men and women and for workers with high and low levels of education, since the public sector premium may not be constant for different levels of education. I also estimate regressions that include all workers and regressions that include only workers employed in the formal sector. This yields eighteen regressions for each survey, or a total of 1,060 regressions for the full sample of sixty surveys. The estimates and standard errors of γ are reported in tables B1 through B3 in appendix B.

There are three possible problems with OLS estimation of equation 1. First, the equation assumes that workers employed in the public and private sectors have the same structure of returns (that is, the vector β is the same for public and private sector workers). Second, estimations of equation 1 exclude workers who do not participate in the labor market. Since the decision to participate in the labor market depends on the comparison between the market wage and the individual's own reservation wage, the decision to participate may be endogenous and induce a selectivity bias. Finally, if the sorting of workers between public and private sectors is nonrandom, OLS estimations will be biased. Van der Gaag and Vijverberg show that selection bias is an important issue in the estimation of public-private wage differentials.[25] In particular, controlling for selection bias lowers the estimated size of the public sector premium and in some cases leads to the finding that public sector workers are paid less than their private sector counterparts.

The first of these problems can be easily addressed by decomposing the differences between public and private wages using the method first developed by Blinder and Oaxaca.[26] The second issue, in turn, can be resolved following Heckman's method, provided appropriate instruments can be identified. There is no easy fix for the third problem, however. Unless one wants to identify a switching regression model using the nonlinearity of the selection equation, solving this problem requires the identification of an instrument for the choice of the sector of employment, that is, a variable that is correlated with the sector of employment but that does not affect a

24. The matrix **X** includes the following variables: experience (defined as age minus years of education minus 6); experience squared; five education dummies (some primary school, completed primary school, some secondary school, completed secondary school, and more than secondary school; no schooling is the excluded dummy); eight dummies for the sector of occupation (mining; manufacture; construction; water and electricity; retail, restaurant and hotel; transport and telecommunications; financial services; and other services); and a dummy for the area of residence (differentiating urban and rural areas).

25. Van der Gaag and Vijverberg (1988).

26. Blinder (1973); Oaxaca (1973).

worker's earnings. Panizza and Qiang address this problem by testing the robustness of OLS estimation of equation 1 to different assumptions on the correlation between the wage and selection equations.[27] The present analysis focuses on OLS estimation of equation 1. In most cases, the results are robust to controlling for selectivity bias and relaxing the assumption that workers employed in the public and private sectors have the same returns (see appendix A).

Results

Table 3 summarizes the results of more than 1,000 regressions and gives an overview of the unweighted regional averages. When all public sector workers are considered, the public sector premium averages 14 percent (4 percent if workers employed in the informal sector are excluded from the sample). A small public sector premium is also found for male workers (approximately 4 percent), together with a much higher public sector premium for female workers (approximately 27 percent). In the case of male workers, most of the public-private wage differential is due to the low wages paid in the informal sector (which is private). The public sector premium for male workers completely disappears when only formal sector workers are considered, and the public sector premium for female workers decreases to approximately 10 percent. Another common feature of the public-private wage differential in most Latin American countries is that the premium tends to be higher for both male and female workers with low levels of education. Table 3 reports an average 3 percent public sector penalty for male workers with high levels of education employed in the formal sector, which falls to zero when workers employed in the informal sector are included in the analysis. The public sector premium for male workers with low levels of education employed in the formal sector is also 3 percent, but it increases to 8 percent when informal sector workers are included. The difference between workers with high education and workers with low education is stronger in the sample of female workers. In this case, the public sector premium for workers with low education employed in the formal sector is approximately 15 percent (38 percent if informal sector workers are included), while the public sector premium for formal sector workers with high education is approximately 4 percent (13 percent if the informal sector is included).

27. Panizza and Qiang (1999).

TABLE 3. Public Sector Premium in Latin America[a]

		All sectors		Formal sector only		
Period	*All workers*	*High education*	*Low education*	*All workers*	*High education*	*Low education*
All workers						
All years	0.144	0.057	0.162	0.046	0.013	0.058
1993–1999	0.143	0.058	0.162	0.027	0.003	0.035
Men						
All years	0.043	–0.007	0.080	0.001	–0.026	0.026
1993–1999	0.043	–0.007	0.080	–0.012	–0.031	0.005
Women						
All years	0.267	0.126	0.383	0.112	0.052	0.183
1993–1999	0.258	0.126	0.369	0.080	0.030	0.145

Source: Author's calculations.
a. Unweighted regional averages. Negative values indicate a public sector penalty.

The public sector premium also varies across age groups. Table 4 reports regional averages for public-private wage differential for three groups of workers: workers aged 20–34, workers aged 35–55, and workers aged 56–65.[28] In the case of men, the public sector premium increases (or the penalty decreases) with age. The full sample of all available surveys yields a 1.4 percent average penalty for young public sector workers and 1.6 percent average premium for older public sector workers. The situation is somewhat different in the case of women. The public sector premium reaches a minimum for middle-aged women and a maximum for older women.

The averages presented in tables 3 and 4 hide the fact that many countries have a significant public sector premium while a much smaller number of countries have a significant public sector penalty. In fact, if I only consider countries that are characterized by a significant (positive or negative) difference between public and private sector wages, the average public sector premium is approximately 7 percent for male workers (0.5 percent if only formal sector workers are included) and approximately 28 percent for female workers (14 percent if only formal sector workers are considered).

Tables B1 through B3 in appendix B show that the seventeen countries studied exhibit large cross-country differences. In the case of male workers, for instance, Bolivia, the Dominican Republic, Honduras, and Panama

28. The full set of regression results is available on request.

TABLE 4. Public Sector Premium in Latin America, by Age Group[a]

		Age group		
Period	*All workers*	*20–34*	*35–55*	*56–65*
Men (formal sector)				
All years	0.001	–0.014	0.008	0.016
1993–1999	–0.012	–0.024	–0.006	–0.005
Women (formal sector)				
All years	0.112	0.110	0.090	0.130
1993–1999	0.080	0.090	0.060	0.110

Source: Author's calculations.
a. Unweighted regional averages. Negative values indicate a public sector penalty.

are characterized by large public sector penalties, which reach 30 percent in the Dominican Republic. At the same time, Colombia, Costa Rica, Ecuador, and El Salvador feature large public sector premiums. Brazil is a strange case, with a large positive premium when only formal sector workers are included in the analysis and a coefficient that is not statistically significant when all workers are considered. In the case of female workers, I find a significant public sector premium in Chile, Colombia, Costa Rica, Ecuador, Guatemala, Honduras, Mexico, El Salvador, and Venezuela and a significant public sector penalty in Nicaragua and Uruguay.

A comparison of the evolution of the public sector premium and public sector employment indicates that during the last decade, some countries (Bolivia, Costa Rica, Honduras, and Uruguay) substantially reduced the size of the public sector by cutting both public sector wages and public sector employment. In Panama and Venezuela, the reduction of public sector employment was accompanied by an increase in relative public sector wages.

Most of the OLS results discussed in this section are robust to controlling for selection in the labor market and selection between sectors of employment (see the appendix A). However, the results of the sensitivity analysis suggest that the OLS estimations could be problematic for Bolivia, Brazil, and Honduras. In addition, the data used to compute the public-private wage differentials reported in tables B1 through B3 do not include nonwage benefits. Given the evidence that nonwage benefits tend to be higher in the public sector, my estimations of public-private wage differentials are likely to underestimate the true rent enjoyed by public sector workers.[29]

Another problem with the estimations of this section is that they are based on a common econometric specification that does not take into account what could be important differences in the structure of the labor market across countries. While this exercise is useful for providing an overview of public-private wage differentials in Latin America, the results need to be corroborated by in-depth country studies.[30] An important aspect that is not considered in this paper is the role of state-owned enterprises, since the data do not distinguish civil servants from employees of state-owned enterprises. In-depth country studies would probably facilitate making this distinction—or at least generate educated guesses based on the area of residence or on privatization patterns. In their study of privatization in Mexico, La Porta and Lopez de Silanes find that the privatization process was followed by an increase in average wages paid by former state-owned enterprises.[31] In addition, contrary to what is suggested in this paper, the wages of blue-collar workers rose more than those of white-collar workers. These results seem to indicate that remuneration of civil servants follows a very different pattern from the remuneration of employees of state-owned enterprises.

In-depth country studies would also reveal the characteristics of different classes of public sector workers. For instance, it would be interesting to compare public-private wage differentials in sectors in which the pub-

29. Quinn (1982) finds that public sector employees in the United States receive pension contributions that are 30 to 50 percent greater than the pension contributions paid by private employers. It has also been estimated that one-third of the raw differential in total labor cost between public and private sector employees can be attributed to nonwage benefits. Brunelli and Cox (1992) claim that only if a federal employee's starting salary is 33.7 percent below the salary of a comparable private employee would there be no excess rent for the public employee. Since wage regressions underestimate the real difference between public and private sector compensation, indirect methods are often used to measure the rent captured by public sector employees. One option consists of estimating whether there is a "queue" for public sector jobs; another involves comparing the quit rates of workers in each sector. Longer queues or lower quit rates in the public sector would indicate that there is a rent to be collected from working in the public sector. Brunelli and Cox (1992) claim that the average tenure of U.S. nonmilitary federal government workers is approximately three times that of private sector workers. The authors use this assertion as indirect evidence that federal workers are overpaid. Long (1982) also finds strong evidence of low quit rates in the public sector. Ippolito (1987) presents a dissenting view, claiming that once one controls for the different pension systems, the turnover for U.S. federal employees is not significantly different from the turnover in the private sector.

30. See, for instance, Amarante (2001).

31. La Porta and Lopez de Silanes (1999).

lic sector is in direct competition with the private sector, such as education and health care.

Many theories that aim at explaining public-private wage differentials focus on the role of public sector unions.[32] Although explaining the differentials is not the goal of this paper, it would be interesting to look at whether public-private wage differentials are significantly correlated with the degree of unionization of public sector workers. Unfortunately, I was not able to find data that differentiate public sector union density from private sector union density. I therefore focus on the correlation between public-private wage differentials and total union density as a percentage of nonagricultural labor force.[33] The data show that the seventeen Latin American countries had an average union density of 14 percent in the mid-1990s. Guatemala and Honduras had the lowest union density (less than 5 percent of the nonagricultural labor force), while Brazil and Mexico had the highest union density (over 31 percent of the nonagricultural labor force).

Table 5 measures the correlation between public-private wage differentials and union density in the mid-1990s, distinguishing between all formal sector workers, formal sector workers with low education, and formal sector workers with high education. The results for the full sample show no significant correlation between public-private wage differential and unionization. If the two countries with the highest degree of unionization (Brazil and Mexico) are dropped, the data yield a significant negative correlation between unionization and the wage differential for both all formal workers and the subset of formal workers with high education. If most cross-country variation in union density comes from differences in private sector union density, the results of table 5 can be reconciled with theories that postulate a positive relationship between public-private wage differential and the relative degree of unionization in the public and private sectors. However, a formal test of this prediction would require disaggregate data for public and private sector union density.

32. See, for instance, Holmlund (1993).
33. The data on union density are from ILO (1998).

TABLE 5. Unionization and Public-Private Wage Differentials[a]

Measure	All observations	Excluding Brazil and Mexico
PRA[b]	0.019	–0.543**
	(0.940)	(0.036)
PRL[c]	–0.056	–0.346
	(0.833)	(0.207)
PRH[d]	–0.027	–0.697*
	(0.919)	(0.004)

Source: Author's calculations.
* Statistically significant at the 1 percent level.
** Statistically significant at the 5 percent level.
a. *P* values in parentheses.
b. Public-private wage differential for all formal sector workers.
c. Public-private wage differential for formal sector workers with low education.
d. Public-private wage differential for formal sector workers with high education.

Public-Private Wage Differentials and the Quality of the Public Sector

Given that institutional quality and economic growth are clearly related and that most developing countries are plagued by corrupt and inefficient public sectors, any policy that might increase the efficiency of the public sector would have large economic returns.[34] Theoretical models that study the incentives for individuals to supply effort or not to engage in activities that go against the interest of the principal (for instance, accepting bribes) show that high wages, in combination with monitoring, can be used to provide the right set of incentives.[35] This argument is often used as the basis for claiming that higher wages for civil servants are a necessary condition for reducing public sector corruption and improving the efficiency of the public sector.[36] However, public sector wages could be endogenous and may depend on the public sector's ability to monitor its agents. Acemoglu and Verdier, for instance, suggest that "the possibility of corruption is likely to increase the size of the government and public sector wages, as compared to the case where corruption was not possible."[37] High public sector wages could thus be a signal of the government's inability to mon-

34. See, for instance, Mauro (1995); Knack and Keefer (1995).
35. See, for instance, Becker and Stigler (1974).
36. A representative example of this view is the following: "The [Singapore] government believed that an efficient bureaucratic system is the one in which officers are well-paid so the temptation to resort to bribes would be reduced.'' Rahman (1986, p. 151), quoted in Van Rijckeghem and Weder (2001, p. 307).
37. Acemoglu and Verdier (2000, p. 209).

itor its agents and, more generally, of low institutional quality. In a case study of the corruption crackdown in the city of Buenos Aires, di Tella and Schargrodsky argue that higher wages reduce corruption only when combined with good audit policies. They conclude that "exclusive emphasis on wage raises may be misplaced and . . . carrots and sticks should be viewed as complementary tools in fighting corruption."[38]

Although it is commonly believed that the remuneration of public sector employees should be positively correlated with the efficiency of the public sector, there is surprisingly little empirical evidence on this topic. Until very recently, the idea that high public sector wages lead to high bureaucratic quality was mostly based on "one-observation econometrics" (namely, the case of Singapore).[39] The main limitation to conducting cross-country empirical research on the relation between public sector wages and the efficiency (or corruption) of the bureaucracy is the lack of an internationally comparable data set on public sector wages. In the last few years, however, Rauch and Evans, as well as Van Rijckeghem and Weder, have begun to build such a data set and test the relation between the remuneration of public sector employees and the quality of the public sector.[40]

Rather than examining the whole public sector, Rauch and Evans collected data on top officials (defined as holding the top 500 positions) in core economic agencies, usually the Ministry of Finance and the Central Bank.[41] Their data set covers thirty-five low- and middle-income countries and focuses on three indices of the bureaucratic structure of core economic agencies. The first index measures whether recruitment is meritocratic at the entry level. The second measures the possibility of internal promotions and career stability. The third measures the level and trend of the ratio between the pay of top civil servants and the pay of private sector workers with similar competence levels. Rauch and Evans's main finding is a strong positive correlation between meritocratic recruitment and bureaucratic quality, but no robust correlation between bureaucratic quality and the other two measures of bureaucratic structure.

38. Di Tella and Schargrodsky (2000, p. 20).

39. In contrast, a large body of empirical work studies the correlation between institutional quality (or corruption) and industrial structure and openness (Sachs and Warner, 1995), origin of the legal code (La Porta and others, 1999; Chong and Zanforlin, 2000), electoral rules (Panizza, 1999; Persson, Tabellini, and Trebbi, 2001), and mortality rates faced by European colonialists (Acemoglu, Johnson, and Robinson, 2000).

40. Rauch and Evans (2000); Van Rijckeghem and Weder (1997).

41. Rauch and Evans (2000).

Van Rijckeghem and Weder, in turn, look at the correlation between corruption and government wages (using the ratio between average government wages and average manufacturing wages) for thirty-one developing and low-income OECD countries over the period 1982–94. While they find a robust negative correlation between public-private wage differentials and the level of corruption, their estimations imply that if a developing country hopes to reduce corruption to the level of high-income OECD countries, it should set its public sector wages at a level that ranges between two and eight times the level of manufacturing wages in the private sector.[42] Van Rijckeghem and Weder suggest, however, that the effect of wages on corruption could be amplified by their indirect effect through rule of law and the quality of the bureaucracy and that eradication of corruption could be achieved with public sector wages that are substantially lower than what is indicated by the paper's estimations.

Treisman attempts to estimate the relation between corruption and public sector wages using a data set on public sector employment assembled by Schiavo-Campo, de Tommaso, and Mukherjee; he does not find any significant effect of public sector employment on corruption. Using the same data set, La Porta and others report a positive correlation between public sector wages and corruption.[43]

However, the measures of public-private wage differentials used by Rauch and Evans, Van Rijckeghem and Weder, Treisman, and La Porta and others contain a number of problems. Because Rauch and Evans focus on top employees in top economic agencies, their results cannot be extended to the whole wage structure of the public sector (and, in fact, this is not their objective). Van Rijckeghem and Weder, in contrast, look at the whole public sector, but their measure of the public-private wage differential is problematic because cross-country differences in the ratio of government wages relative to manufacturing wages could stem from cross-country differences in skill content rather than to differences in the relative remuneration in the two sectors. The data set used by both Treisman and La Porta and others is more problematic still in that it focuses on the ratio of the average government wage to GDP, which is heavily influenced by a coun-

42. Van Rijckeghem and Weder (1997). They use this last finding to suggest that the data are more supportive of Akerlof and Yellen's (1990) fair wage effort hypothesis than of the traditional model presented by Becker and Stigler (1974).

43. Treisman (2000); Schiavo-Campo, de Tommaso, and Mukherjee (1997); La Porta and others (1999).

try's level of economic development—which, in turn, is highly correlated with the quality of the public sector.

The data set on public-private wage differentials assembled in this paper addresses the two issues described above. The differentials are computed using nationally representative surveys, such that they capture public-private wage differentials for the whole public sector, and they are obtained from Mincerian wage regressions in which the public-private differentials are computed holding skills constant. Furthermore, household-level data allow the computing of disaggregate figures for public-private wage differentials for workers with different levels of education and the testing of interesting hypotheses on the correlation between the degree of meritocracy and the overall quality of the public sector.

This section analyzes the correlation between the public-private wage differential and two indices of the quality of the public sector. The first measure of quality is the International Country Risk Guide (ICRG) index of bureaucratic quality (BURQ). The ICRG defines bureaucratic quality as high when the bureaucracy tends to be autonomous from political pressure and when it has an established mechanism for recruitment and training. The second measure is the ICRG index of corruption (CORR), which aims at measuring both bureaucratic corruption and potential corruption within the political system. CORR captures corruption in a very indirect way, however. One of the main indicators used to build this index is the length of time a government has been in power continuously. Specifically, the ICRG guide defines the measure as follows:

> In assessing the corruption risk, therefore, we look first at how long a government has been in power continuously. In the case of one-party state or non-elected government, corruption in the form of patronage and nepotism is an essential prerequisite and it is therefore corrupt, to a great or lesser degree, from its inception. In the case of a democratic government, it has been our experience, almost without exception, that things begin to go wrong after an elected government has been in office for more than two consecutive terms, that is eight to ten years. On that basis, the highest risk ratings tend to signify an accountable democracy whose government has been in office for less than five years. An intermediate rating often indicates a country whose government has been in office for more than ten years and where a large number of officials are appointed rather than elected. The lowest ratings are usually given to one-party states and autarchies.[44]

44. Sealy (1999, p. 13).

Reverse causality is a serious issue. A number of mechanisms could yield a positive causal link between the length of time a government has been in power continuously and public sector wages. For instance, public sector workers are often the main supporters of autocratic regimes that have been in power for a long time, and the regimes may reward this loyalty with generous remuneration. The shortcomings of this index are illustrated by the fact that in 1995 Singapore received the same grade as Mozambique, Syria, and Turkey. Even with these caveats, I include CORR in my analysis because this index has been used in previous work and is, therefore, useful in comparing the results of this paper with previous work.

The paper explores the correlation between the quality of the public sector and five measures of the public-private wage structure. These measures include three indices of public-private wage differentials: one computed for all formal sector workers (PRA) (the coefficients of column 6 of table B3); one computed for all formal sector workers with low education (PRL) (the coefficients of column 8 of table B3); and one computed for all formal sector workers with high education (PRH) (the coefficients of column 7 of table B3).[45] My fourth measure of the public-private wage structure is the ratio between the third and second indices: RHL = (PRH + 1) / (PRL + 1). This can be interpreted as a measure of the relative public-private wage differentials for workers with high and low skills.[46] Finally, I measure the relative meritocracy in the public and private sectors by using the ratio between the return to education in the public sector and the return to education in the private sector (REL). This measure could not be obtained from the regressions of equation 1 because the equation assumes a nonlinear correlation between the log of wages and education. REL was instead calculated by assuming a linear relationship between education and the log of wages. Formally, the first step was to estimate the following regression separately for public and private sector workers:

$$w_i = \mathbf{X}_i\boldsymbol{\beta} + \mathrm{EDUC}_i\xi + u_i, \tag{2}$$

45. I always use formal sector workers because the sample that includes informal sector workers exhibits a high correlation between the public sector premium and the size of the informal sector, which, in turn, is negatively correlated with public sector quality.

46. Define WPUH and WPUL as public sector wages for workers with high and low skills, respectively, and WPRH and WPRL as private sector wages for workers with high and low skills, respectively. Then PRH = log(WPUH) – log(WPRH) and PRL = log(WPUL) – log(WPRL). Therefore, RHL = (PRH + 1)/(PRL + 1) = (WPUH/WPRH)/(WPUL/WPRL).

where w_i is the log of hourly earnings, $\mathbf{X}_i$ is a matrix of productivity-related characteristics (namely, experience, experience squared, eight dummies for the sector of occupation, and a dummy for the area of residence), and $EDUC_i$ is a variable measuring years of education. Next, REL was obtained by dividing ξ (the coefficient attached to $EDUC_i$) in the regression that included only public sector workers by the corresponding coefficient in the regression that included only private sector workers.

Given that my sample includes sixty surveys from seventeen countries over the 1981–99 period, it would be tempting to treat each survey as a single observation and run a fixed effects estimation over this sample of sixty surveys. This method has two serious drawbacks, however. First, the panel is highly unbalanced: seven countries have one or two observations (Chile, the Dominican Republic, Ecuador, Guatemala, Mexico, Nicaragua, and Paraguay), four countries have three or four observations (Bolivia, El Salvador, Panama, and Peru), four countries have five or six observations (Brazil, Colombia, Honduras, and Uruguay), and two countries have seven observations (Costa Rica and Venezuela). Treating each survey as one observation would give excessive weight to countries with many surveys. For instance, the weight of Costa Rica would be six times that of Chile.

Second, in most countries the surveys were taken at a very close distance from one another. Although the surveys span the 1981–99 period, approximately 60 percent of the surveys (thirty-six out of sixty surveys) were collected in 1993–98. This is problematic because the dependent variables used in the regressions (BURQ and CORR) have limited within-country variability (see table 6). Of the thirteen countries for which I have more than one observation, BURQ and CORR are constant in seven. The limited within-country variability of the ICRG indices can also be shown by regressing the indices on a set of country dummies. These country dummies alone (that is, the fixed effects) explain 86 percent of the variability of BURQ and 89 percent of the variability of CORR. Using fixed effects estimations would therefore eliminate 90 percent of the variability of the data.

Since standard fixed effects are problematic, I average the data for the pre- and post-1993 periods and use random effects regressions to estimate the correlation between public-private wage differentials and the efficiency of the public sector. The data used in the regressions are reported in table B4. The results of the sixth column of the table (REL) may appear

TABLE 6. Within-Country Variation of the International Country Risk Guide (ICRG) Indices

Country	*BURQ*[a]	*CORR*[b]	*Country*	*BURQ*[a]	*CORR*[b]
Bolivia	0.38	0.23	Honduras	0.00	0.00
Brazil	0.00	0.16	Panama	0.35	0.00
Chile	0.20	0.20	Paraguay	0.00	0.00
Colombia	0.00	0.21	Peru	0.00	0.00
Costa Rica	0.06	0.00	Uruguay	0.40	0.00
Ecuador	0.00	0.20	Venezuela	0.07	0.14
El Salvador	0.00	0.00	Cross-country	0.28	0.30

Source: Author's calculations.
a. Bureaucratic quality.
b. Corruption.

to contradict the results of columns 4 and 5 (PRL and PRH). In the cases of Chile, Colombia, Costa Rica, and Panama, PRL is higher than PRH (indicating that the public-private wage differential decreases with education) but REL is more than one (indicating that returns to education are higher in the public sector). This is due to the fact that the econometric specification used to estimate REL is different from the one used to estimate PRL and PRH. Specifically, REL was estimated by assuming a linear relationship between education and the log of wages, whereas PRL and PRH were estimated by allowing for a nonlinear relationship between education and the log of wages.

While the main drawback of fixed effects estimations is that they do not make use of cross-country variability, random effects estimations are problematic because they cannot keep track of all unobserved country-specific characteristics. If these country-specific characteristics are correlated with the regressors, random effects yield biased estimations. However, most variables that have been found to be correlated with institutional quality, like religion and the origin of the legal code, are constant across Latin America, and they are thus implicitly controlled for.[47] Furthermore, tables 7 and 8 indicate that in most cases the Hausmann test does not reject the hypothesis of $E(u_i, \mathbf{X}_i) = 0$ and, therefore, does not reject the appropriateness of the random effects model. The costs of using fixed effects estimations thus seem to outweigh those of random effect estimations.[48]

47. See, for instance, La Porta and others (1999).

48. Fixed effects estimations generate results that are very close to those presented in table 7. However, the results of these fixed effects estimations are based on regressions with eight degrees of freedom!

TABLE 7. Efficiency of the Bureaucracy and Public-Private Wage Differentials, Random Effects Estimations[a]

Explanatory variable	*(1)*	*(2)*	*(3)*	*(4)*	*(5)*	*(6)*	*(7)*	*(8)*	*(9)*
PRA	0.300	0.656							
	(0.35)	(0.80)							
PRL			−0.373	−0.001			−0.720		
			(0.61)	(0.02)			(1.17)		
PRH					1.633	1.552	2.042		
					(1.78)***	(1.82)***	(2.17)**		
RHL								0.998	
								(1.70)***	
REL									1.032
									(2.20)**
GDP		0.854		0.823		0.785	0.698	0.682	0.767
		(2.43)**		(2.27)**		(2.33)**	(2.00)**	(1.91)***	(2.28)**
CONS	2.510	−4.345	2.550	−4.065	2.504	−3.780	−3.040	−3.900	−4.721
	(11.52)*	(1.54)	(11.34)*	(1.39)	(12.39)*	(1.40)	(1.08)	(1.38)	(1.72)***
Summary statistic									
R^2	0.24	0.36	0.14	0.27	0.27	0.44	0.39	0.25	0.35
Hausmann test	4.01	2.23	6.48	5.18	0.22	0.03	3.53	3.66	0.51
P value	0.045	0.33	0.01	0.08	0.64	0.98	0.32	0.16	0.77
No. observations	27	27	27	27	27	27	27	27	27

Source: Author's calculations.

* Statistically significant at the 1 percent level.

** Statistically significant at the 5 percent level.

*** Statistically significant at the 10 percent level.

a. Dependent variable is the ICRG index of bureaucratic quality. Explanatory variables are the public-private wage differential for all formal sector workers (PRA); the public-private wage differential for formal sector workers with low education (PRL); the public-private wage differential for formal sector workers with high education (PRH); the ratio between the third and second indices [RHL = (PRH + 1) / (PRL + 1)]; the ratio between the return to education in the public sector and the return to education in the private sector (REL); the log of GDP per capita; and a constant (CONS). Absolute value of *t* statistics in parentheses.

TABLE 8. Corruption and Public-Private Wage Differentials, Random Effects Estimations[a]

Explanatory variable	*(1)*	*(2)*	*(3)*	*(4)*	*(5)*	*(6)*	*(7)*	*(8)*	*(9)*
PRA	0.670	0.701							
	(0.81)	(0.83)							
PRL			0.059	0.086			−0.616		
			(0.09)	(0.13)			(1.01)		
PRH					1.639	1.652	2.135		
					(1.87)***	(1.85)***	(2.18)**		
RHL								0.649	
								(1.00)	
REL									0.366
									(0.65)
GDP		0.096		0.074		0.059	−0.017	−0.003	0.057
		(0.25)		(0.19)		(0.16)	(0.04)	(0.01)	(0.15)
CONS	2.987	2.214	3.017	2.425	3.001	2.531	3.172	2.417	2.173
	(14.29)*	(0.72)	(14.21)*	(0.77)	(14.08)*	(0.84)	(1.03)	(0.78)	(0.69)
Summary statistic									
R^2	0.10	0.12	0.07	0.07	0.08	0.08	0.05	0.001	0.06
Hausmann test	0.24	0.17	0.63	0.67	1.3	1.76	3.53	1.58	0.03
P value	0.62	0.92	0.43	0.73	0.25	0.42	0.32	0.45	0.86
No. observations	27	27	27	27	27	27	27	27	27

Source: Author's calculations.

* Statistically significant at the 1 percent level.

** Statistically significant at the 5 percent level.

*** Statistically significant at the 10 percent level.

a. Dependent variable is the ICRG index of corruption. Explanatory variables are the public-private wage differential for all formal sector workers (PRA); the public-private wage differential for formal sector workers with low education (PRL); the public-private wage differential for formal sector workers with high education (PRH); the ratio between the third and second indices [RHL = (PRH + 1) / (PRL + 1)]; the ratio between the return to education in the public sector and the return to education in the private sector (REL); the log of GDP per capita; and a constant (CONS). Absolute value of *t* statistics in parentheses.

Table 7 illustrates that there is no significant correlation between public-private wage differentials for all workers (PRA) and bureaucratic quality. The coefficient is positive but not statistically significant. Columns 3, 4, 5, and 6 reveal an interesting pattern, however. The public-private wage differential for workers with low education (PRL) has a negative coefficient, and the public-private wage differential for workers with high education (PRH) has a positive and statistically significant coefficient. This result holds if PRL and PRH are included in the same regression (column 7). In this case, PRH is statistically significant and significantly different from PRL (the hypothesis that the PRL = PRH is rejected with a *p* value of 0.04). Column 8 similarly shows that there is a positive and statistically significant correlation between bureaucratic quality and the ratio between high- and low-education public-private wage differentials (RHL). Finally, the last column indicates that there is a positive and statistically significant correlation between bureaucratic quality and the ratio between the return to education in the public sector and the return to education in the private sector (REL). The point estimations of table 7 imply that a one standard deviation change in RHL or REL is associated with a 0.2 to 0.3 standard deviation change in bureaucratic quality.

Although the limited number of observations prevents a more detailed analysis of the correlation between public-private wage differentials and bureaucratic quality, the results of table 7 are suggestive and, given the small sample size, particularly strong. The only control included in the regressions of table 7 is the log of GDP per capita, but the results are robust to the inclusion of other factors that have been found to be associated with institutional quality. In particular, the results are robust to the inclusion of a measure of openness (see Sachs and Warner) and to the inclusion of a variable measuring a country's average number of years of education.[49]

Table 7 thus suggests that while there is no correlation between the average public-private wage differential and bureaucratic quality, there is a positive correlation between bureaucratic quality and the differential between the relative remuneration of workers with low education and workers with high education. The regressions also indicate a significant correlation between the relative return to education in the public and private sectors and bureaucratic quality. These results corroborate Rauch and

49. Sachs and Warner (1995). The results are available on request.

Evans's finding that meritocracy at the entry level is strongly associated with the quality of the public sector, as well as Katz and Krueger's finding that in the United States, the wage compression of the public sector has lowered the quality of white-collar workers in the public sector (as measured by standardized tests).[50]

Since most of the literature that studies the relationship between public sector wages and the quality of the public sector focuses on corruption, I now examine the correlation between the structure of the public-private wage differential and the ICRG index of corruption (see table 8). As in the case of bureaucratic quality, I do not find any significant correlation between corruption and the average public-private wage differential. The correlation between PRL and CORR is positive but not significant, and the correlation between PRH and CORR is positive and statistically significant. When PRL and PRH are entered in the same regression (column 7), the coefficient attached to the first variable switches sign and becomes negative, while the coefficient attached to PRH remains positive and statistically significant. The coefficients of PRL and PRH are significantly different form each other (the hypothesis that PRL = PRH is rejected with a p-value of 0.04). Both RHL and RAL have positive coefficients, but neither coefficient is statistically significant.

It is fair to conclude that even though the results for CORR follow the pattern of the results for BURQ, the former are much weaker and, in most cases, indicate no significant relationship between corruption and the structure of public-private wage differentials. There are two possible explanations for this finding. On the one hand, the mechanisms that determine bureaucratic quality may be different from those that determine corruption, and the structure of the public-private wage differential may have a stronger effect on the former. On the other hand, the main variable used to evaluate corruption (that is, the length of time a government has been in power continuously) may not be a good proxy for a country's actual level of corruption. The fact that in 1995 Singapore was ranked as having the same level of corruption as many developing countries, including Albania, Madagascar, Mozambique, Syria, and Zimbabwe, seems to lend support to the latter hypothesis.

50. Rauch and Evans (2000); Katz and Krueger (1991).

Conclusion

It is often claimed that public sector workers are underpaid, especially in developing countries, and that there is a close association between public sector wages and either the efficiency or the level of corruption in the public sector. This paper builds a data set of public-private wage differentials for a sample of seventeen Latin American countries covering 88 percent of the population in the region. Three stylized facts are revealed: on average, Latin American countries are characterized by a public sector premium; on average, the public sector premium tends to be higher for women than for men; and on average, the public sector premium is higher for workers with low education. In fact, workers with high education may suffer a public sector penalty.

The paper then uses the data set on public-private wage differentials to examine whether there is a significant correlation between the structure of public-private wage differential and the quality of the public sector. The results indicate that in Latin America, bureaucratic quality and average public-private wage differential are not significantly correlated. There is a significant correlation, however, between bureaucratic quality and the degree of meritocracy of the public sector. The paper finds weak evidence of a significant relationship between the structure of public-private wage differential and corruption.

Several caveats apply to these results. First, they are based on a small sample of countries and therefore only relate to Latin America. Second, the limited number of observations does not support an in-depth robustness analysis. It would be interesting to test whether these results are robust to the inclusion of other low- and middle-income countries. Third, the public-private wage differentials were estimated using a common econometric specification for all seventeen countries in the sample. This is clearly a rough first pass at the data. In-depth country studies are needed to generate more precise estimations of the structure of public-private wage differentials.

Even with these caveats, a series of policy lessons can be drawn from the results. Public sector wage inequality tends to be lower than private sector wage inequality. Public sector employees are also characterized by a lower wage gender gap, given that the public sector premium is higher for women than for men.[51] Notwithstanding this lower wage inequality, there is no consensus on the redistributive effects of public sector employment. Rodrik points out that public sector employment is less affected by the business cycle than is private sector employment and can thus play an important role in reducing income risk.[52] The Inter-American Development Bank, in contrast, illustrates that public sector employment is often a regressive aspect of public spending.[53] This paper suggests that even if public sector employment is not regressive, lower wage inequality within the public sector may come at a cost. In particular, the paper finds a positive correlation between the degree of meritocracy of the public sector and bureaucratic quality, which implies that low wage inequality in the public sector is associated with low bureaucratic quality. The regressions presented in table 7 do not specify the direction of the causality. However, if causality goes from public sector wages to bureaucratic quality, then increasing average public sector wages may not be the most efficient policy for improving the quality of the public sector. In particular, Van Rijckeghem and Weder estimate that quasi-eradication of corruption would require average public sector wages to be two to eight times higher than average manufacturing wages.[54] The results of this paper suggest that the same outcome could be reached by increasing the degree of meritocracy within the public sector. This is likely to carry a lower cost than an across-the-board increase in public sector wages. It would therefore free resources to implement inequality reduction through more progressive components of public spending, such as health and education, and to develop income security policies that would affect all social groups, not just those employed in the public sector.[55]

Appendix A: Checking the Robustness of the OLS Estimations

As mentioned in the text, the estimation of equation 1 is econometrically problematic for at least three reasons. First, it models the sectoral effects

51. See Panizza and Qiang (1999).

52. Rodrik (1997). Alesina, Danninger, and Rostagno (1999) show that in the case of Italy, public sector employment plays an important role in redistributing resources from the richer northern part of the country to the poorer southern part of the country.

53. IDB (1999).

54. Van Rijckeghem and Weder (1997).

55. IDB (1999).

as intercept effects and holds returns to other productivity-related characteristics and job attributes equal across sectors. Second, it excludes workers who do not participate in the formal labor market, which may induce a selectivity bias. Third, if the choice of the sector of employment (public versus private) is nonrandom, OLS estimations of equation 1 will yield biased coefficients. This appendix demonstrates that controlling for these factors does not alter the basic results of the paper. The first two issues are addressed by decomposing the public-private wage differentials using the decomposition suggested by Blinder and Oaxaca and by running Heckit estimations for female workers. The third issue is dealt with by estimating public-private wage differentials under different assumptions for the correlation between selection and wage equations. To make the results readable, the sensitivity analysis focuses on formal sector workers and employs a subset of surveys measured in the mid-1990s.

To relax the assumption of equal return in the public and private labor market, table A1 decomposes public-private average wage differentials using the method first developed by Blinder and Oaxaca.[56] This method allows separating the difference in average worker characteristics between sectors from the difference in the returns to worker characteristics between sectors. Since I have two sectors (public and private) and two groups of workers (men and women), I need to estimate four sets of equations. Formally, I estimate the following two equations for both men and women:

$$w_i^g = \mathbf{X}_i^g \boldsymbol{\beta}^g + \varepsilon_i^g \tag{3}$$

and

$$w_i^p \mathbf{X}_i^p \boldsymbol{\beta}^p + \varepsilon_i^p, \tag{4}$$

where w_i^g and w_i^p are the logs of hourly earnings in the public and private sectors, respectively, and $\mathbf{X}_i^g$ and $\mathbf{X}_i^p$ are vectors of productivity-related characteristics. Next, I use $\hat{\boldsymbol{\beta}}^g$ and $\hat{\boldsymbol{\beta}}^p$ to decompose the difference in average earnings in the two sectors into a component for the difference in average worker characteristics and a component for the difference in the returns to worker characteristics. Formally,

$$\overline{w}^g - \overline{w}^p = \left(\overline{\mathbf{X}}^g - \overline{\mathbf{X}}^p\right)' \hat{\boldsymbol{\beta}}^* + \left[\overline{\mathbf{X}}^{g'}\left(\hat{\boldsymbol{\beta}}^g - \hat{\boldsymbol{\beta}}^*\right) - \overline{\mathbf{X}}^{p'}\left(\hat{\boldsymbol{\beta}}^p - \hat{\boldsymbol{\beta}}^*\right)\right], \tag{5}$$

56. Blinder (1973); Oaxaca (1973).

TABLE A1. Public-Private Wage Differentials Using OLS, Heckit, and Oaxaca Decomposition, Formal Sector Workers Only

	Men				Women				
		Oaxaca Decomposition					Oaxaca Decomposition		
Country and year of survey	OLS	Total	Endow-ment	Premium	OLS	Heckit	Total	Endow-ment	Premium
Bolivia 1997	−0.19*	1.06	1.14	0.93	−0.20*	−0.19**	1.30	1.41	0.91
Brazil 1995	0.28*	2.88	2.72	1.06	0.19*	0.05	2.91	2.77	1.05
Chile 1996	−0.02	1.48	1.51	0.98	−0.07	0.13*	1.27	1.35	0.94
Colombia 1997	0.21*	2.19	1.92	1.14	0.25*	0.29*	2.15	1.85	1.16
Costa Rica 1995	0.13*	1.63	1.54	1.06	0.14*	0.19*	1.92	1.83	1.04
Dominican Rep. 1996	−0.32*	0.77	0.94	0.82	−0.01	0.25*	1.12	1.16	0.97
Ecuador 1995	0.11***	1.43	1.36	1.06	0.01	−0.02	1.30	1.33	0.98
El Salvador 1995	0.18**	1.84	1.71	1.08	0.27*	0.28*	2.17	1.94	1.12
Guatemala 1998	−0.14	1.45	1.47	0.98	0.24*	0.27*	1.95	1.74	1.12
Honduras 1996	0.01	1.47	1.45	1.01	0.28*	0.20*	1.73	1.55	1.12
Mexico 1994	0.01	1.43	1.36	1.06	0.11**	0.22*	1.30	1.33	0.98
Nicaragua 1993	−0.09***	1.23	1.31	0.94	−0.13***	−0.37**	1.08	1.20	0.90
Panama 1995	−0.24*	1.37	1.51	0.91	0.02	−0.02	1.48	1.48	1.01
Paraguay 1995	0.05	1.52	1.49	1.01	0.14***	0.03	1.44	1.36	1.05
Peru 1997	−0.07	1.32	1.38	0.95	−0.01	−0.08	1.57	1.56	0.95
Uruguay 1995	−0.08*	1.05	1.10	0.96	−0.08*	−0.07**	1.27	1.34	0.96
Venezuela 1995	−0.04***	1.1	1.12	0.97	−0.003	−0.02	1.06	1.06	1.00

Source: Author's calculations.
* Statistically significant at the 1 percent level.
** Statistically significant at the 5 percent level.
*** Statistically significant at the 10 percent level.

where $\hat{\boldsymbol{\beta}}^*$ measures the return to worker characteristics that would exist in the absence of an unequal rate of return between the two sectors. The decomposition presented in table A1 was carried out by setting $\hat{\boldsymbol{\beta}}^*$ equal to the return to worker characteristics for the pooled sample of public and private sector employees (similar results are obtained by setting $\hat{\boldsymbol{\beta}}^* = \hat{\boldsymbol{\beta}}^g$ or $\hat{\boldsymbol{\beta}}^* = \hat{\boldsymbol{\beta}}^p$). In the setting of equation 5, the first term on the right-hand side measures the difference in average wages stemming from differences in worker characteristics, and the second term on the right-hand side measures the difference in return to these characteristics, that is, the public sector. Finally, the left-hand side of equation 5 measures the total raw average wage differential. Table A1 reports the results of the OLS and Heckit estimations as $\overline{w}^g - \overline{w}^p$ and the results for the Oaxaca decomposition as W^g/W^p (where W is the level, and not the log, of earnings). In the case of Bolivia, for example, the –0.19 coefficient in the OLS column indicates a public sector penalty of 19 percent and the 0.93 coefficient in the premium column indicates a public sector penalty of 7 percent.

The results of the Oaxaca decomposition go in the same direction as the results of the OLS estimations. Although the results obtained with the two methods are not identical, in only one case (namely, female workers in Mexico) does the OLS regression suggest a significant public sector premium while the Oaxaca decomposition suggests a small public sector penalty.

Another source of concern with the OLS estimations discussed in the text is that participation in the labor market can be endogenous, which may represent a serious problem in the case of female workers. Table A1 therefore reports public-private wage differentials obtained by estimating a wage equation with the method developed by Heckman.[57] The results of the Heckit estimation are reassuring. While the coefficients are not identical to the OLS estimations, the two methods yield contradictory results for only two countries (Chile and Dominican Republic). In both cases, OLS estimations indicate a nonsignificant public sector penalty, and the Heckit estimations suggest a significant public sector premium.

The Oaxaca decomposition has the advantage of not imposing the same vector, $\boldsymbol{\beta}$, on public and private sector workers, but it is still affected by a possible source of selection bias in that a worker's choice between work-

57. Heckman (1979). I use the following variables to identify the wage equation: family status dummies (single, married, divorced, widower), other family income, and number of children under six years of age.

ing in the private sector and working in the public sector could be nonrandom. If this is the case, it is important to model the selection process and discuss the consequences of nonrandom sorting between the two sectors. Stelcner, van der Gaag, and Vijverberg model the selection process with the following equation:[58]

$$I_i^* = \mathbf{Z}_i\gamma + u_i, \tag{6}$$

where I_i^* takes a value of 1 if the worker belongs to the public sector and 0 if the worker belongs to the private sector. If the choice between the two sectors is endogenous, OLS estimation of equations 3 and 4 provides unbiased estimates of $\boldsymbol{\beta}$ only if $E(u, \varepsilon^g) = E(u, \varepsilon^p) = 0$.[59]

If it were possible to find a set of variables that affect the selection process but are not a determinant of a worker's remuneration, it would be possible to obtain unbiased estimates of $\boldsymbol{\beta}$ by running full-information maximum likelihood estimations of a model that includes equations 3, 4, and 6. However, while I could identify a set of variables to use in modeling the decision of whether to participate in the labor market, I do not have an instrument with which to model the choice between the two sectors. I therefore follow Panizza and Qiang and proceed as follows.[60] I start by assuming that selection bias is not an important issue and estimate equations 3 and 4 with OLS. I then reestimate the model under different assumptions for the correlation between the selection and wage equations. In many cases, the results are robust to a wide range of values for this correlation. Formally, I estimate equations 7 and 8 under different assumptions for $\rho^g = E(u, \varepsilon^g)/\sigma_u\sigma_{\varepsilon^g}$ and $\rho^p = E(u, \varepsilon^p)/\sigma_u\sigma_{\varepsilon^p}$:

$$w_i^g - \rho^g\sigma_{\varepsilon^g}\lambda_i = \mathbf{X}_i\boldsymbol{\beta} + v_i^g \text{ and} \tag{7}$$

$$w_i^p - \rho^p\sigma_{\varepsilon^p}\Lambda_i = \mathbf{X}_i\boldsymbol{\beta} + v_i^p, \text{ with} \tag{8}$$

$$\lambda_i = \frac{f(-\mathbf{Z}_i\gamma)}{1 - F(-\mathbf{Z}_i\gamma)} \tag{9}$$

$$\Lambda_i = \frac{f(\mathbf{Z}_i\gamma)}{F(-\mathbf{Z}_i\gamma)}. \tag{10}$$

58. Stelcner, van der Gaag, and Vijverberg (1989).
59. See van der Gaag and Vijverberg (1988).
60. Panizza and Qiang (1999).

In estimating equations 7 and 8, I set σ_{ε^g} and σ_{ε^p} equal to the standard deviations of the residuals of equations 3 and 4, respectively. While $\rho^g \in (-1, 1)$ and $\rho^p \in (-1, 1)$, I assume that there is positive selection in both sectors and set $\rho^g \in (0, 1)$ and $\rho^p \in (-1, 0)$.Table A2 presents the results of the simulation and compares them with the benchmark case of no selection ($\rho^g = \rho^p = 0$).[61]

For male workers, using different assumptions for the correlation between the selection and wage equations yields very different results in three countries (Bolivia, Brazil, and Honduras). The possible role of selection seems to be more important in the sample of female workers. In this case, I find seven countries (Brazil, Costa Rica, Ecuador, Guatemala, Honduras, Mexico, and Paraguay) for which different assumptions for the correlation between the selection and wage equations are associated with sign swings in the estimations of the public-private wage differential.

61. Table A2 reports the results for ρ^g and ρ^p in the (0, 0.5) and (–0.5, 0) ranges. Results for the (0, 1) and (–1, 0) ranges are available on request.

TABLE A2. Sensitivity Analysis for Public-Private Wage Differentials

	$\rho^g = 0$			$\rho^g = 0.25$			$\rho^g = 0.50$		
Country	$\rho^p = 0$	$\rho^p = -0.25$	$\rho^p = -0.50$	$\rho^p = 0$	$\rho^p = -0.25$	$\rho^p = -0.50$	$\rho^p = 0$	$\rho^p = -0.25$	$\rho^p = -0.50$
Male workers									
Bolivia	0.93	1.66	1.57	0.96	1.71	1.62	0.98	1.76	1.67
Brazil	1.06	1.05	1.03	0.92	0.91	0.89	0.93	0.92	0.91
Chile	0.98	0.97	0.95	1.01	0.99	0.98	1.04	1.03	1.02
Colombia	1.14	1.12	1.09	1.20	1.18	1.15	1.26	1.24	1.21
Costa Rica	1.06	1.03	1.02	1.08	1.06	1.04	1.11	1.08	1.07
Ecuador	1.06	1.02	0.99	1.13	1.09	1.06	1.21	1.17	1.14
Guatemala	0.98	0.95	0.93	1.01	0.98	0.96	1.04	1.01	1.0
Honduras	1.01	0.95	0.93	1.10	1.04	1.02	1.13	1.07	1.04
Mexico	1.06	1.04	1.03	1.09	1.08	1.06	1.13	1.11	1.10
Nicaragua	0.94	0.92	0.89	0.97	0.94	0.92	1.00	0.97	0.94
Panama	0.91	0.86	0.84	0.89	0.85	0.83	0.91	0.86	0.84
Paraguay	1.01	1.06	1.03	1.05	1.10	1.08	1.10	1.15	1.12
Peru	0.95	0.93	0.91	0.98	0.95	0.94	1.01	0.98	0.96
Dominican Rep.	0.82	0.84	0.82	0.84	0.85	0.84	0.85	0.86	0.86
El Salvador	1.08	1.07	1.06	1.10	1.08	1.07	1.11	1.10	1.09
Uruguay	0.97	0.96	0.94	0.98	0.97	0.96	0.99	0.98	0.97
Venezuela	0.97	0.95	0.94	1.00	0.98	0.97	1.03	1.02	1.00

Female workers									
Bolivia	0.91	0.84	0.77	0.94	0.86	0.80	0.96	0.89	0.82
Brazil	1.05	0.97	0.94	1.02	0.94	0.91	1.03	0.95	0.92
Chile	0.94	0.92	0.90	0.97	0.95	0.92	1.00	0.98	0.95
Colombia	1.16	1.13	1.10	1.22	1.19	1.62	1.27	1.24	1.22
Costa Rica	1.04	0.84	0.82	1.08	0.88	0.86	1.12	0.91	0.89
Ecuador	0.98	0.91	0.84	1.02	0.95	0.88	1.07	0.99	0.92
Guatemala	1.12	0.88	0.81	1.19	0.93	0.85	1.26	0.98	0.91
Honduras	1.12	0.83	0.80	1.14	0.85	0.81	1.20	0.89	0.85
Mexico	0.98	0.95	0.92	1.02	0.99	0.96	1.06	1.03	1.00
Nicaragua	0.90	0.89	0.89	0.91	0.90	0.89	0.92	0.91	0.90
Panama	1.01	0.99	0.98	0.97	0.96	0.94	0.99	0.97	0.96
Paraguay	1.05	0.90	0.87	1.12	0.95	0.93	1.14	0.98	0.95
Peru	0.95	0.92	0.89	0.99	0.96	0.93	1.03	1.00	0.96
Dominican Rep.	0.97	0.97	0.93	0.99	0.99	0.95	1.01	1.01	0.97
El Salvador	1.12	1.11	1.09	1.14	1.14	1.12	1.17	1.16	1.14
Uruguay	0.96	0.93	0.91	0.98	0.95	0.93	0.99	0.98	0.96
Venezuela	1.00	0.97	0.94	1.02	0.99	0.96	1.04	1.01	0.98

Source: Author's calculations.

Appendix B: Supplemental Tables

TABLE B1. Public-Private Wage Differential, Male Workers[a]

		All workers			*Formal sector workers only*		
Country	*Year*	*All*	*High education*	*Low education*	*All*	*High education*	*Low education*
Bolivia	1990	–0.286*	–0.373*	–0.223*	–0.123**	–0.334*	0.074
		(0.041)	(0.056)	(0.058)	(0.053)	(0.077)	(0.068)
	1993	–0.107**	–0.135**	–0.232*	–0.101***	–0.149**	–0.119
		(0.048)	(0.062)	(0.069)	(0.055)	(0.071)	(0.077)
	1996	–0.153*	–0.221*	–0.090	–0.127**	–0.138***	–0.057
		(0.052)	(0.068)	(0.084)	(0.060)	(0.076)	(0.099)
	1997	–0.147**	–0.042	–0.317*	–0.187**	–0.101	–0.371**
		(0.065)	(0.071)	(0.114)	(0.075)	(0.078)	(0.151)
Brazil	1992	0.146*	0.139*	0.152*	0.597*	0.615*	0.570*
		(0.018)	(0.025)	(0.026)	(0.031)	(0.059)	(0.038)
	1993	–0.052*	0.087*	–0.145*	0.264*	0.605*	0.098*
		(0.019)	(0.028)	(0.027)	(0.031)	(0.062)	(0.037)
	1995	0.003	0.000	0.033	0.282*	0.285*	0.261*
		(0.018)	(0.024)	(0.026)	(0.032)	(0.055)	(0.040)
	1996	–0.016	–0.036	0.042***	0.301*	0.362*	0.256*
		(0.017)	(0.024)	(0.025)	(0.029)	(0.055)	(0.037)
	1997	0.022	0.017	0.064*	0.382*	0.445*	0.338*
		(0.017)	(0.022)	(0.024)	(0.027)	(0.044)	(0.036)
Chile	1987	0.004	–0.064**	–0.010	0.061**	–0.019	0.063***
		(0.023)	(0.030)	(0.032)	(0.024)	(0.031)	(0.032)
	1996	–0.053	–0.072	–0.110*	–0.018	–0.064	–0.028
		(0.043)	(0.045)	(0.036)	(0.043)	(0.043)	(0.037)
Colombia	1990	0.141*	0.108*	0.176*	0.194*	0.146*	0.222*
		(0.021)	(0.027)	(0.030)	(0.023)	(0.029)	(0.032)
	1991	0.062*	0.010	0.134*	0.103*	0.051	0.157*
		(0.021)	(0.031)	(0.029)	(0.023)	(0.033)	(0.032)
	1993	0.140*	0.082*	0.201*	0.185*	0.140*	0.212*
		(0.018)	(0.025)	(0.026)	(0.020)	(0.027)	(0.028)
	1997	0.245*	0.190*	0.339*	0.210*	0.181*	0.209*
		(0.023)	(0.028)	(0.040)	(0.024)	(0.030)	(0.038)
	1998	0.207*	0.134*	0.286*	0.172*	0.108*	0.222*
		(0.025)	(0.032)	(0.038)	(0.025)	(0.033)	(0.038)
	1999	0.184*	0.165*	0.234*	0.225*	0.193*	0.245*
		(0.033)	(0.042)	(0.057)	(0.036)	(0.046)	(0.059)
Costa Rica	1983	0.129*	0.029	0.186*	0.077*	–0.102**	0.182*
		(0.024)	(0.043)	(0.028)	(0.030)	(0.045)	(0.037)
	1985	0.093*	0.068	0.110*	0.049***	–0.023	0.085**
		(0.024)	(0.046)	(0.028)	(0.029)	(0.053)	(0.034)
	1987	0.163*	0.096***	0.185*	0.265*	0.283*	0.227*
		(0.027)	(0.051)	(0.030)	(0.033)	(0.067)	(0.035)

TABLE B1. Continued

		All workers			*Formal sector workers only*		
Country	*Year*	*All*	*High education*	*Low education*	*All*	*High education*	*Low education*
	1991	0.294*	0.368*	0.246*	0.236*	0.288*	0.194*
		(0.028)	(0.051)	(0.032)	(0.034)	(0.061)	(0.037)
	1993	0.173*	0.146*	0.179*	0.129*	0.058	0.178*
		(0.030)	(0.049)	(0.036)	(0.034)	(0.051)	(0.037)
	1995	0.175*	0.170*	0.119*	0.129*	0.107**	0.107*
		(0.028)	(0.043)	(0.037)	(0.031)	(0.048)	(0.039)
	1997	0.167*	0.169*	0.132*	0.117*	0.102***	0.115*
		(0.028)	(0.048)	(0.034)	(0.033)	(0.053)	(0.039)
Dominican Rep.	1996	–0.368*	–0.277*	–0.435*	–0.319*	–0.228*	–0.388*
		(0.034)	(0.059)	(0.041)	(0.040)	(0.067)	(0.049)
Ecuador	1995	0.225*	0.073	0.373*	0.106***	–0.041	0.285*
		(0.053)	(0.075)	(0.069)	(0.064)	(0.090)	(0.091)
	1998	0.376*	0.322*	0.489*	0.271*	0.248*	0.334*
		(0.057)	(0.077)	(0.080)	(0.067)	(0.088)	(0.092)
El Salvador	1995	0.276*	0.041	0.438*	0.178*	0.043	0.289*
		(0.036)	(0.054)	(0.044)	(0.038)	(0.057)	(0.048)
	1997	0.260*	0.232*	0.301*	0.243*	0.242*	0.272*
		(0.034)	(0.049)	(0.045)	(0.036)	(0.052)	(0.047)
	1998	0.264*	0.121***	0.376*	0.175*	0.096	0.234*
		(0.041)	(0.065)	(0.045)	(0.041)	(0.067)	(0.044)
Guatemala	1998	–0.045	0.060	–0.169	–0.136	0.041	–0.285**
		(0.082)	(0.089)	(0.125)	(0.087)	(0.089)	(0.133)
Honduras	1989	0.136*	0.131**	0.143*	0.039	–0.008	0.063
		(0.029)	(0.052)	(0.035)	(0.034)	(0.054)	(0.042)
	1992	0.104**	0.015	0.151*	0.057	–0.011	0.099
		(0.047)	(0.070)	(0.057)	(0.053)	(0.068)	(0.071)
	1996	0.017	0.083	–0.042	0.011	0.013	–0.008
		(0.047)	(0.074)	(0.062)	(0.050)	(0.082)	(0.065)
	1997	–0.091	–0.004	–0.138***	–0.325*	–0.160***	–0.440*
		(0.059)	(0.083)	(0.078)	(0.072)	(0.083)	(0.109)
	1998	–0.098	0.003	–0.245*	–0.286*	–0.047	–0.594*
		(0.061)	(0.083)	(0.089)	(0.081)	(0.103)	(0.122)
Mexico	1994	0.110*	0.070	0.109*	0.007	–0.035	0.010
		(0.032)	(0.055)	(0.038)	(0.039)	(0.071)	(0.046)
Nicaragua	1993	–0.022	–0.191**	0.039	–0.090***	–0.181***	–0.060
		(0.048)	(0.090)	(0.057)	(0.053)	(0.096)	(0.063)
Panama	1979	0.052**	–0.008	0.079***	0.052**	–0.008	0.079***
		(0.021)	(0.045)	(0.023)	(0.021)	(0.045)	(0.023)
	1991	0.128***	–0.027	0.249***	–0.194***	–0.249***	–0.136***
		(0.033)	(0.049)	(0.044)	(0.035)	(0.050)	(0.048)
	1995	0.095***	–0.154***	0.271***	–0.239***	–0.295***	–0.195***
		(0.032)	(0.046)	(0.041)	(0.037)	(0.050)	(0.055)

(*continued*)

TABLE B1. Continued

		All workers			*Formal sector workers only*		
Country	*Year*	*All*	*High education*	*Low education*	*All*	*High education*	*Low education*
	1997	0.181***	–0.078	0.413***	–0.125***	–0.230***	0.028
		(0.036)	(0.053)	(0.046)	(0.040)	(0.056)	(0.055)
Paraguay	1998	0.116	–0.175	0.580***	–0.093	–0.419***	0.326
		(0.162)	(0.195)	(0.191)	(0.202)	(0.216)	(0.250)
	1995	0.099**	–0.089	0.283***	0.045	–0.094	0.173**
		(0.049)	(0.070)	(0.065)	(0.06)	(0.088)	(0.082)
Peru	1985	0.181***	0.121***	0.222***	0.051	–0.087	0.176***
		(0.052)	(0.069)	(0.078)	(0.076)	(0.100)	(0.103)
	1994	–0.072	–0.151**	0.085	–0.121**	–0.174***	–0.002
		(0.052)	(0.063)	(0.084)	(0.055)	(0.065)	(0.097)
	1996	0.040	–0.003	0.032	–0.028	–0.076***	–0.013
		(0.029)	(0.034)	(0.049)	(0.038)	(0.044)	(0.065)
	1997	–0.001	–0.004	0.113	–0.074	–0.053	–0.006
		(0.057)	(0.061)	(0.118)	(0.058)	(0.060)	(0.127)
Uruguay	1981	0.111	–0.005	0.029	–0.070**	–0.071	–0.088**
		(0.020)	(0.040)	(0.025)	(0.030)	(0.060)	(0.036)
	1989	0.030***	–0.036	0.069***	–0.120***	–0.170***	–0.090***
		(0.017)	(0.030)	(0.020)	(0.020)	(0.040)	(0.026)
	1992	–0.002	–0.140***	0.080***	–0.130***	–0.220***	–0.090***
		(0.020)	(0.040)	(0.025)	(0.020)	(0.040)	(0.028)
	1995	0.065***	0.004	0.130***	–0.079***	–0.120***	–0.027
		(0.016)	(0.026)	(0.020)	(0.019)	(0.029)	(0.024)
	1997	0.168***	–0.005	0.287***	–0.006	–0.130***	0.110***
		(0.017)	(0.027)	(0.020)	(0.018)	(0.029)	(0.020)
Venezuela	1981	–0.054***	0.019	–0.075***	–0.016	0.034	–0.027***
		(0.010)	(0.024)	(0.011)	(0.013)	(0.029)	(0.014)
	1983	–0.050***	0.035***	–0.074***	–0.081***	–0.013	–0.098***
		(0.008)	(0.019)	(0.008)	(0.009)	(0.021)	(0.010)
	1986	–0.007	0.022	–0.027***	–0.064***	–0.042***	–0.082***
		(0.009)	(0.020)	(0.010)	(0.011)	(0.023)	(0.012)
	1989	–0.013	–0.011	–0.013	–0.060***	–0.048**	–0.063***
		(0.010)	(0.020)	(0.011)	(0.011)	(0.021)	(0.012)
	1993	–0.098***	–0.055***	–0.128***	–0.096***	–0.071***	–0.101***
		(0.011)	(0.020)	(0.012)	(0.013)	(0.023)	(0.014)
	1995	0.050**	0.040	0.061**	–0.041***	–0.055	–0.031
		(0.022)	(0.035)	(0.029)	(0.025)	(0.040)	(0.031)
	1997	0.168***	0.132***	0.167***	0.081**	0.064	0.109***
		(0.027)	(0.042)	(0.032)	(0.032)	(0.050)	(0.036)

Source: Author's calculations.

* Statistically significant at the 1 percent level.

** Statistically significant at the 5 percent level.

*** Statistically significant at the 10 percent level.

a. Standard errors are in parentheses.

TABLE B2. Public-Private Wage Differential, Female Workers[a]

		All workers			*Formal sector workers only*		
Country	*Year*	*All*	*High education*	*Low education*	*All*	*High education*	*Low education*
Bolivia	1990	–0.231*	–0.350*	–0.109	–0.029	–0.139	0.116
		(0.056)	(0.075)	(0.083)	(0.077)	(0.098)	(0.123)
	1993	0.197*	0.027	0.190**	–0.092	–0.108	0.061
		(0.060)	(0.077)	(0.090)	(0.085)	(0.088)	(0.143)
	1996	0.131**	0.091	0.219***	0.077	–0.014	0.283***
		(0.065)	(0.072)	(0.129)	(0.078)	(0.084)	(0.147)
	1997	–0.046	0.010	–0.180	–0.203**	–0.132	–0.458***
		(0.073)	(0.080)	(0.183)	(0.081)	(0.082)	(0.264)
Brazil	1992	–0.049*	0.017	–0.154*	0.289*	0.362*	0.196*
		(0.017)	(0.022)	(0.026)	(0.037)	(0.051)	(0.051)
	1993	–0.095*	–0.070*	–0.124*	0.294*	0.390*	0.172*
		(0.017)	(0.022)	(0.026)	(0.040)	(0.056)	(0.052)
	1995	–0.080*	–0.078*	–0.095*	0.190*	0.249*	0.105**
		(0.015)	(0.020)	(0.024)	(0.036)	(0.050)	(0.052)
	1996	–0.084*	–0.090*	–0.090*	0.232*	0.281*	0.150*
		(0.015)	(0.019)	(0.025)	(0.032)	(0.046)	(0.043)
	1997	–0.075*	–0.084*	–0.074*	0.201*	0.217*	0.128*
		(0.015)	(0.018)	(0.023)	(0.032)	(0.042)	(0.047)
Chile	1987	0.242*	0.073**	0.438*	0.276*	0.104*	0.498*
		(0.026)	(0.030)	(0.050)	(0.026)	(0.029)	(0.051)
	1996	0.091**	–0.049	0.202*	–0.071***	–0.102**	–0.114
		(0.038)	(0.038)	(0.059)	(0.042)	(0.040)	(0.101)
Colombia	1990	0.289*	0.196*	0.369*	0.332*	0.229*	0.413*
		(0.022)	(0.026)	(0.037)	(0.023)	(0.026)	(0.038)
	1991	0.124*	0.119*	0.099***	0.123*	0.124*	0.065
		(0.029)	(0.030)	(0.054)	(0.031)	(0.031)	(0.057)
	1993	0.359*	0.206*	0.469*	0.389*	0.221*	0.54*
		(0.02)	(0.025)	(0.036)	(0.021)	(0.025)	(0.039)
	1997	0.352*	0.230*	0.461*	0.254*	0.211*	0.189*
		(0.025)	(0.026)	(0.049)	(0.022)	(0.024)	(0.045)
	1998	0.289*	0.231*	0.281*	0.242*	0.219*	0.123
		(0.038)	(0.028)	(0.104)	(0.037)	(0.029)	(0.101)
	1999	0.180*	0.128*	0.180*	0.204*	0.139*	0.222*
		(0.034)	(0.038)	(0.064)	(0.035)	(0.039)	(0.068)
Costa Rica	1983	0.631*	0.249*	0.871*	0.118**	0.067	0.207***
		(0.045)	(0.069)	(0.056)	(0.056)	(0.067)	(0.116)
	1985	0.549*	0.303*	0.717*	0.112**	0.062	0.266*
		(0.040)	(0.057)	(0.054)	(0.046)	(0.049)	(0.099)
	1987	0.552*	0.187**	0.713*	0.752*	0.491*	0.76*
		(0.046)	(0.075)	(0.054)	(0.059)	(0.149)	(0.06)
	1991	0.498*	0.275*	0.596*	0.179*	0.102	0.287*
		(0.045)	(0.066)	(0.062)	(0.055)	(0.071)	(0.09)

(*continued*)

TABLE B2. Continued

		All workers			Formal sector workers only		
Country	*Year*	*All*	*High education*	*Low education*	*All*	*High education*	*Low education*
	1993	0.333*	0.106***	0.407*	0.040	–0.028	0.103
		(0.042)	(0.057)	(0.057)	(0.050)	(0.057)	(0.090)
	1995	0.378*	0.249*	0.361*	0.137*	0.111**	0.193**
		(0.038)	(0.049)	(0.057)	(0.046)	(0.049)	(0.08)
	1997	0.349*	0.212*	0.397*	0.124*	0.017	0.276*
		(0.036)	(0.049)	(0.052)	(0.045)	(0.049)	(0.080)
Dominican Rep.	1996	0.231*	0.022	0.370*	–0.008	0.008	–0.034
		(0.048)	(0.062)	(0.071)	(0.056)	(0.068)	(0.085)
Ecuador	1995	0.173*	0.157**	0.310**	0.008	0.041	–0.157
		(0.064)	(0.078)	(0.130)	(0.082)	(0.093)	(0.180)
	1998	0.345*	0.375*	0.545*	0.262*	0.114	0.808*
		(0.077)	(0.087)	(0.158)	(0.101)	(0.098)	(0.204)
El Salvador	1995	0.617*	0.347*	0.809*	0.265*	0.248*	0.277**
		(0.044)	(0.058)	(0.087)	(0.051)	(0.056)	(0.118)
	1997	0.631*	0.406*	0.797*	0.667*	0.324*	0.982*
		(0.043)	(0.060)	(0.066)	(0.051)	(0.070)	(0.073)
	1998	0.762*	0.434*	0.871*	0.314*	0.295*	0.308*
		(0.046)	(0.063)	(0.080)	(0.056)	(0.070)	(0.097)
Guatemala	1998	0.396*	0.380*	0.443*	0.242**	0.269*	0.135
		(0.082)	(0.094)	(0.136)	(0.108)	(0.099)	(0.197)
Honduras	1989	0.787*	0.178*	1.103*	0.185*	0.051	0.391*
		(0.043)	(0.054)	(0.058)	(0.049)	(0.050)	(0.095)
	1992	0.667*	0.269*	0.835*	0.236*	0.232*	0.167
		(0.063)	(0.071)	(0.109)	(0.067)	(0.063)	(0.165)
	1996	0.591*	0.268*	0.697*	0.279*	0.209*	0.377*
		(0.050)	(0.063)	(0.071)	(0.055)	(0.065)	(0.096)
	1997	0.473*	0.098	0.737*	0.060	–0.010	0.189***
		(0.056)	(0.065)	(0.083)	(0.059)	(0.064)	(0.108)
	1998	0.491*	0.215**	0.621*	0.099	0.031	0.174
		(0.062)	(0.087)	(0.103)	(0.075)	(0.071)	(0.165)
Mexico	1994	0.233*	0.123**	0.283*	0.110**	0.079	0.147**
		(0.043)	(0.056)	(0.063)	(0.048)	(0.060)	(0.070)
Nicaragua	1993	0.022	–0.144	0.103	–0.127***	–0.382*	0.019
		(0.056)	(0.093)	(0.065)	(0.069)	(0.092)	(0.093)
Panama	1979	0.468*	0.067	0.626*	0.468*	0.067	0.626*
		(0.032)	(0.058)	(0.039)	(0.032)	(0.058)	(0.039)
	1991	0.509*	0.338*	0.626*	–0.011	–0.002	–0.025
		(0.035)	(0.052)	(0.046)	(0.046)	(0.052)	(0.082)
	1995	0.502*	0.347*	0.667*	0.022	0.022	–0.048
		(0.032)	(0.046)	(0.045)	(0.043)	(0.045)	(0.119)
	1997	0.473*	0.306*	0.638*	0.057	0.030	0.077
		(0.032)	(0.044)	(0.045)	(0.044)	(0.047)	(0.078)

TABLE B2. Continued

		All workers			Formal sector workers only		
Country	*Year*	*All*	*High education*	*Low education*	*All*	*High education*	*Low education*
Paraguay	1998	0.242	0.171	1.092*	−0.101	−0.142	0.324
		(0.151)	(0.184)	(0.282)	(0.162)	(0.184)	(0.456)
	1995	0.316*	0.158**	0.613*	0.138***	0.035	0.405*
		(0.058)	(0.077)	(0.092)	(0.083)	(0.099)	(0.125)
Peru	1985	0.369*	0.419*	0.254	0.191**	0.209**	0.120
		(0.075)	(0.085)	(0.196)	(0.096)	(0.093)	(0.308)
	1994	0.076	−0.037	0.112	−0.077	−0.102	0.109
		(0.062)	(0.067)	(0.193)	(0.075)	(0.072)	(0.321)
	1996	0.155*	0.090***	−0.030	0.106	−0.003	0.493***
		(0.043)	(0.050)	(0.210)	(0.070)	(0.068)	(0.271)
	1997	0.188*	0.072	0.091	−0.006	−0.098	0.277
		(0.059)	(0.064)	(0.141)	(0.073)	(0.070)	(0.208)
Uruguay	1981	0.400*	0.230*	0.490*	0.230**	0.160	0.280*
		(0.027)	(0.048)	(0.034)	(0.040)	(0.066)	(0.058)
	1989	0.030	0.006	0.048	−0.190*	−0.150*	−0.210*
		(0.020)	(0.040)	(0.030)	(0.030)	(0.040)	(0.040)
	1992	0.080*	−0.080**	0.228*	−0.130*	0.200*	−0.040
		(0.026)	(0.038)	(0.036)	(0.028)	(0.038)	(0.040)
	1995	0.130*	−0.017	0.279*	−0.080*	−0.139*	−0.010
		(0.018)	(0.026)	(0.025)	(0.020)	(0.027)	(0.030)
	1997	0.130*	0.0003	0.268*	−0.120*	−0.160*	−0.060***
		(0.018)	(0.020)	(0.030)	(0.020)	(0.030)	(0.034)
Venezuela	1981	0.260*	0.092*	0.270*	0.225*	0.123*	0.228*
		(0.015)	(0.033)	(0.016)	(0.023)	(0.044)	(0.026)
	1983	0.240*	0.101*	0.279*	0.060*	0.058**	0.074*
		(0.012)	(0.027)	(0.013)	(0.015)	(0.028)	(0.017)
	1986	0.246*	0.105*	0.291*	0.072*	0.087*	0.052*
		(0.012)	(0.023)	(0.014)	(0.015)	(0.025)	(0.018)
	1989	0.327*	0.070*	0.447*	0.048*	0.039***	0.054*
		(0.014)	(0.023)	(0.017)	(0.015)	(0.020)	(0.020)
	1993	0.152*	−0.065*	0.278*	−0.067*	−0.119*	0.001
		(0.013)	(0.020)	(0.015)	(0.016)	(0.021)	(0.020)
	1995	0.277*	0.053	0.431*	−0.003	−0.037	0.044
		(0.025)	(0.036)	(0.033)	(0.029)	(0.038)	(0.043)
	1997	0.366*	0.276*	0.389*	0.141*	0.173*	0.022
		(0.027)	(0.041)	(0.035)	(0.036)	(0.047)	(0.044)

Source: Author's calculations.
* Statistically significant at the 1 percent level.
** Statistically significant at the 5 percent level.
*** Statistically significant at the 10 percent level.
a. Standard errors are in parentheses.

TABLE B3. Public-Private Wage Differential, All Workers[a]

		All workers			*Formal sector workers only*		
Country	*Year*	*All*	*High education*	*Low education*	*All*	*High education*	*Low education*
Bolivia	1990	–0.269*	–0.375*	–0.238*	–0.092**	–0.273*	0.044
		(0.033)	(0.045)	(0.046)	(0.044)	(0.062)	(0.060)
	1993	0.064***	–0.068	–0.039	–0.094**	–0.140**	–0.117***
		(0.037)	(0.049)	(0.054)	(0.047)	(0.057)	(0.069)
	1994	–0.047	–0.095***	–0.012	–0.049	–0.092	0.039
		(0.041)	(0.049)	(0.070)	(0.049)	(0.057)	(0.083)
	1997	–0.103**	–0.009	–0.262*	–0.194*	–0.109***	–0.357*
		(0.049)	(0.054)	(0.096)	(0.055)	(0.056)	(0.129)
Brazil	1992	0.029**	0.064*	–0.017	0.474*	0.480*	0.418*
		(0.012)	(0.016)	(0.018)	(0.024)	(0.039)	(0.030)
	1993	–0.081*	–0.007	–0.152*	0.277*	0.510*	0.101*
		(0.013)	(0.017)	(0.018)	(0.024)	(0.042)	(0.029)
	1994	–0.051*	–0.051*	–0.040**	0.238*	0.267*	0.182*
		(0.012)	(0.015)	(0.017)	(0.024)	(0.037)	(0.032)
	1996	–0.058*	–0.069*	–0.038**	0.269*	0.322*	0.178*
		(0.011)	(0.015)	(0.017)	(0.021)	(0.036)	(0.028)
	1997	–0.037*	–0.045*	–0.020	0.304*	0.331*	0.232*
		(0.011)	(0.014)	(0.016)	(0.020)	(0.031)	(0.028)
Chile	1987	0.103*	0.001	0.091*	0.157*	0.041***	0.171*
		(0.017)	(0.021)	(0.026)	(0.017)	(0.021)	(0.026)
	1994	0.006	–0.071**	–0.009	–0.044	–0.084*	–0.078***
		(0.028)	(0.029)	(0.030)	(0.030)	(0.030)	(0.041)
Colombia	1990	0.218*	0.152*	0.260*	0.275*	0.189*	0.323*
		(0.015)	(0.019)	(0.023)	(0.016)	(0.020)	(0.024)
	1991	0.099*	0.056*	0.091*	0.123*	0.081*	0.103*
		(0.018)	(0.022)	(0.031)	(0.020)	(0.023)	(0.034)
	1993	0.254*	0.143*	0.311*	0.308*	0.182*	0.370*
		(0.014)	(0.017)	(0.021)	(0.015)	(0.018)	(0.022)
	1997	0.305*	0.210*	0.375*	0.250*	0.199*	0.204*
		(0.017)	(0.019)	(0.031)	(0.017)	(0.019)	(0.030)
	1998	0.225*	0.198*	0.161*	0.207*	0.187*	0.092**
		(0.019)	(0.021)	(0.039)	(0.019)	(0.022)	(0.037)
	1999	0.171*	0.142*	0.206*	0.209*	0.162*	0.225*
		(0.024)	(0.028)	(0.042)	(0.025)	(0.029)	(0.044)
Costa Rica	1983	0.316*	0.120*	0.398*	0.090*	–0.035	0.195*
		(0.022)	(0.036)	(0.026)	(0.026)	(0.037)	(0.035)
	1985	0.271*	0.176*	0.289*	0.077*	0.031	0.112*
		(0.021)	(0.035)	(0.025)	(0.025)	(0.037)	(0.033)
	1987	0.326*	0.129*	0.367*	0.470*	0.358*	0.413*
		(0.024)	(0.043)	(0.027)	(0.029)	(0.067)	(0.031)
	1991	0.395*	0.328*	0.369*	0.223*	0.202*	0.210*
		(0.024)	(0.042)	(0.029)	(0.029)	(0.047)	(0.034)

TABLE B3. Continued

		All workers			Formal sector workers only		
Country	*Year*	*All*	*High education*	*Low education*	*All*	*High education*	*Low education*
	1993	0.260*	0.138*	0.264*	0.110*	0.028	0.159*
		(0.024)	(0.038)	(0.030)	(0.028)	(0.039)	(0.035)
	1994	0.268*	0.203*	0.213*	0.132*	0.105*	0.127*
		(0.022)	(0.032)	(0.030)	(0.026)	(0.035)	(0.035)
	1997	0.266*	0.194*	0.229*	0.121*	0.061***	0.147*
		(0.022)	(0.034)	(0.028)	(0.027)	(0.036)	(0.035)
Dominican Rep.	1994	–0.135*	–0.140*	–0.212*	–0.216*	–0.122**	–0.310*
		(0.028)	(0.043)	(0.037)	(0.032)	(0.048)	(0.042)
Ecuador	1994	0.188*	0.110**	0.278*	0.059	–0.004	0.193**
		(0.040)	(0.053)	(0.057)	(0.050)	(0.064)	(0.080)
	1998	0.326*	0.346*	0.387*	0.265*	0.190*	0.490*
		(0.046)	(0.058)	(0.075)	(0.062)	(0.066)	(0.115)
El Salvador	1994	0.466*	0.211*	0.598*	0.214*	0.161*	0.285*
		(0.026)	(0.040)	(0.034)	(0.031)	(0.040)	(0.044)
	1997	0.461*	0.314*	0.486*	0.511*	0.280*	0.606*
		(0.025)	(0.038)	(0.033)	(0.028)	(0.043)	(0.037)
	1998	0.513*	0.275*	0.562*	0.219*	0.194*	0.222*
		(0.030)	(0.046)	(0.038)	(0.034)	(0.049)	(0.042)
Guatemala	1998	0.133**	0.247*	–0.043	0.010	0.187**	–0.202***
		(0.062)	(0.072)	(0.105)	(0.070)	(0.074)	(0.115)
Honduras	1989	0.410*	0.161*	0.442*	0.087*	0.018	0.123*
		(0.025)	(0.038)	(0.032)	(0.028)	(0.037)	(0.038)
	1992	0.389*	0.140*	0.418*	0.129*	0.098**	0.115***
		(0.038)	(0.051)	(0.051)	(0.042)	(0.048)	(0.067)
	1994	0.309*	0.178*	0.244*	0.121*	0.117**	0.085
		(0.035)	(0.049)	(0.051)	(0.038)	(0.052)	(0.059)
	1997	0.251*	0.060	0.284*	–0.132*	–0.077	–0.216*
		(0.041)	(0.052)	(0.060)	(0.048)	(0.051)	(0.081)
	1998	0.238*	0.080	0.170**	–0.112**	0.012	–0.355*
		(0.043)	(0.060)	(0.066)	(0.056)	(0.062)	(0.092)
Mexico	1994	0.175*	0.110*	0.131*	0.052***	0.032	0.036
		(0.025)	(0.039)	(0.032)	(0.031)	(0.048)	(0.039)
Nicaragua	1993	0.000	–0.169**	0.060	–0.101**	–0.261*	–0.040
		(0.036)	(0.067)	(0.042)	(0.042)	(0.070)	(0.052)
Panama	1980	0.260*	0.025	0.323*	0.260*	0.025	0.323*
		(0.019)	(0.037)	(0.021)	(0.019)	(0.037)	(0.021)
	1991	0.345*	0.154*	0.427*	–0.129*	–0.143*	–0.110*
		(0.023)	(0.036)	(0.030)	(0.028)	(0.037)	(0.041)
	1994	0.332*	0.119*	0.436*	–0.130*	–0.141*	–0.171*
		(0.022)	(0.034)	(0.029)	(0.029)	(0.034)	(0.049)
	1997	0.355*	0.137*	0.500*	–0.050***	–0.101*	0.027
		(0.023)	(0.034)	(0.031)	(0.030)	(0.036)	(0.046)

(*continued*)

TABLE B3. Continued

		All workers			*Formal sector workers only*		
Country	*Year*	*All*	*High education*	*Low education*	*All*	*High education*	*Low education*
Paraguay	1995	0.187*	0.029	0.377*	0.074	−0.029	0.226*
		(0.037)	(0.052)	(0.052)	(0.048)	(0.065)	(0.069)
	1994	0.182	−0.032	0.618*	−0.089	−0.259***	0.244
		(0.119)	(0.134)	(0.186)	(0.137)	(0.145)	(0.238)
Peru	1985	0.226*	0.209*	0.188*	0.090	0.015	0.147
		(0.042)	(0.054)	(0.071)	(0.061)	(0.073)	(0.098)
	1994	−0.039	−0.119*	0.045	−0.113**	−0.155*	0.030
		(0.039)	(0.045)	(0.074)	(0.044)	(0.048)	(0.097)
	1996	0.076*	0.032	−0.015	0.030	−0.044	0.106
		(0.024)	(0.029)	(0.049)	(0.035)	(0.038)	(0.072)
	1997	0.066***	0.013	0.099	−0.047	−0.081***	0.083
		(0.040)	(0.044)	(0.080)	(0.046)	(0.046)	(0.115)
Uruguay	1981	0.168*	0.104*	0.155*	0.025	0.022	0.004
		(0.017)	(0.031)	(0.020)	(0.025)	(0.044)	(0.030)
	1989	−0.013	−0.086*	0.018	−0.206*	−0.228*	−0.185*
		(0.015)	(0.026)	(0.018)	(0.018)	(0.031)	(0.022)
	1992	0.046*	−0.104*	0.110*	−0.126*	−0.213*	−0.079*
		(0.016)	(0.027)	(0.020)	(0.018)	(0.029)	(0.023)
	1994	0.095*	−0.016	0.154*	−0.086*	−0.143*	−0.042**
		(0.011)	(0.018)	(0.015)	(0.014)	(0.020)	(0.019)
	1997	0.154*	−0.002	0.257*	−0.048*	−0.150*	0.043**
		(0.012)	(0.019)	(0.016)	(0.014)	(0.020)	(0.020)
Venezuela	1981	0.069*	0.055*	0.046*	0.058*	0.070*	0.043*
		(0.008)	(0.020)	(0.009)	(0.011)	(0.024)	(0.012)
	1983	0.066*	0.065*	0.042*	−0.037*	0.014	−0.062*
		(0.007)	(0.016)	(0.007)	(0.008)	(0.017)	(0.008)
	1986	0.101*	0.059*	0.084*	−0.022**	0.010	−0.057*
		(0.007)	(0.015)	(0.008)	(0.009)	(0.017)	(0.010)
	1988	0.136*	0.029**	0.170*	−0.019**	−0.008	−0.027*
		(0.008)	(0.015)	(0.009)	(0.009)	(0.015)	(0.010)
	1993	0.020**	−0.055*	0.034*	−0.082*	−0.091*	−0.071*
		(0.008)	(0.014)	(0.009)	(0.010)	(0.016)	(0.011)
	1994	0.163*	0.052**	0.216*	−0.025	−0.045	−0.016
		(0.017)	(0.025)	(0.022)	(0.019)	(0.028)	(0.025)
	1997	0.272*	0.207*	0.268*	0.105*	0.118*	0.082*
		(0.019)	(0.029)	(0.023)	(0.024)	(0.035)	(0.028)

Source: Author's calculations.
* Statistically significant at the 1 percent level.
** Statistically significant at the 5 percent level.
*** Statistically significant at the 10 percent level.
a. Standard errors are in parentheses.

TABLE B4. Data Used in the Regressions of Tables 6 and 7

Country	*Period*	*PRA*	*PRL*	*PRH*	*REL*
Bolivia	1980s–1992	−0.092	0.044	−0.273	0.928
	Post-1992	−0.112	−0.145	−0.114	1.163
Brazil	1980s–1992	0.474	0.418	0.480	1.284
	Post-1992	0.272	0.173	0.357	1.414
Chile	1980s–1992	0.157	0.171	0.041	1.221
	Post-1992	−0.044	−0.078	−0.084	1.082
Colombia	1980s–1992	0.199	0.213	0.135	1.054
	Post-1992	0.243	0.223	0.182	1.270
Costa Rica	1980s–1992	0.215	0.232	0.139	1.251
	Post-1992	0.121	0.144	0.064	1.337
Dominican Rep.	1980s–1992	n.a.	n.a.	n.a.	n.a.
	Post-1992	−0.216	−0.310	−0.122	1.249
Ecuador	1980s–1992	n.a.	n.a.	n.a.	n.a.
	Post-1992	0.162	0.341	0.093	0.788
El Salvador	1980s–1992	n.a.	n.a.	n.a.	n.a.
	Post-1992	0.314	0.371	0.211	1.087
Guatemala	1980s–1992	n.a.	n.a.	n.a.	n.a.
	Post-1992	0.010	−0.202	0.187	1.234
Honduras	1980s–1992	0.108	0.119	0.058	0.989
	Post-1992	−0.041	−0.162	0.017	1.474
Mexico	1980s–1992	n.a.	n.a.	n.a.	n.a.
	Post-1992	0.052	0.036	0.032	1.024
Nicaragua	1980s–1992	n.a.	n.a.	n.a.	n.a.
	Post-1992	−0.101	−0.040	−0.261	0.836
Panama	1980s–1992	0.065	0.106	−0.059	1.047
	Post-1992	−0.090	−0.072	−0.121	1.031
Paraguay	1980s–1992	n.a.	n.a.	n.a.	n.a.
	Post-1992	−0.007	0.235	−0.144	0.733
Peru	1980s–1992	0.028	0.074	−0.032	1.037
	Post-1992	−0.044	0.073	−0.093	0.982
Uruguay	1980s–1992	−0.103	−0.087	−0.140	0.837
	Post-1992	−0.067	0.000	−0.146	0.746
Venezuela	1980s–1992	−0.005	−0.026	0.021	1.134
	Post-1992	−0.001	−0.002	−0.006	1.131

Source: Author's calculations.