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Post-Reform Trends in Wage Inequality: The Case of Urban Bolivia

by

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Abstract: This paper seeks to contribute to the ongoing controversy on the distributional effects of structural reforms in developing countries. Applying inequality indices and Fields' (2001) decomposition methodology to Bolivian household survey data of the years 1989 to 1997, we identify recent trends in wage inequality of urban Bolivia. Using a rent-based dual-economy model, we can link these trends to the structural reforms undertaken in Bolivia since 1985.

Key Words: Structural Reforms, Inequality, Dual Economy, Bolivia

JEL Classification: D31, J50, L16, O17

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1. Introduction

Rigidities in product and labor markets are considered to be at the roots of poor economic performance in many developing countries. To restore sustainable economic growth, international organizations such as IMF and World Bank have advocated comprehensive and far-reaching structural reforms in these countries. Critics, however, argue that rising inequality is also part of the deal. The objective of this paper is to assess the validity of this claim for the case of Bolivia.

In recent years, quite a few empirical studies have looked at Bolivian wage inequality. Jemio (2000) observed a rise in wage inequality from 1989 to 1997 as evidenced by a rise in the Gini coefficient and the Theil index. Using different decomposition methodologies, Urquiola (1993) and Fields et al. (1998) found that education was the most important factor explaining the *level* of wage inequality. Andersen (1999) observed that education's explanatory power increased over time. We proceed along this path by decomposing recent *changes* in wage inequality. Furthermore, we set up a theoretical model to link the empirical results to the Bolivian structural reform process.

The paper is structured as follows. In Section 2, we give a summary of the structural reforms undertaken in Bolivia since 1985. Referring to the reform indices of Burki and Perry (1997) and Morley et al. (1999), we show that

Bolivia started structural reforms relatively late. In 1985, all reform indices except the trade and the capital account liberalization index were significantly lower than those of neighbouring Chile. However, due to the fast reform process over the last 15 years, Bolivia has caught up to or even overtaken Chile in most policy areas. Only with respect to labor market reforms, Bolivia still trails substantially behind.

In Sections 3 and 4, we apply inequality indices and Fields' (2001) decomposition methodology to Bolivian household survey data of the years 1989 to 1997, and derive three stylized facts on recent trends in urban wage inequality: (a) We observe that wage inequality increased stronger at the upper tail of the distribution than at the lower tail. (b) We find that education was not only the most important factor explaining the *level* of wage inequality, but also the main contributor to the *rise* in wage inequality. (c) Looking more deeply at the contribution of education to the rise in wage inequality, we discover that a large share of this contribution can be attributed to the increase in the correlation coefficient between schooling years and wages.

In Section 5, we set up a rent-based dual-economy model to explain these empirical results against the background of the Bolivian structural reforms. In line with Saint-Paul (2000), the formal sector is modeled with monopolistically competitive firms, union-firm bargaining and employment protection. In the informal sector, product and labor markets are assumed to be perfectly competitive. We apply the model to study the impact of structural reforms on wages and employment. Following Blanchard and Giavazzi (2001), we introduce the policy measures into the model in a highly abstract fashion by discussing their impact on the model parameters which reflect the market imperfections in the formal sector. Simulating the model numerically, we can replicate the above mentioned stylized facts. Section 6 concludes.

2 Economic Background and Data

2.1 Structural Reforms in Bolivia: An Overview

In the first half of the 1980s, the economic situation in Bolivia was desperate. From 1981 to 1985, the country witnessed five consecutive years of zero or even negative GDP growth; its rate of open unemployment nearly doubled from 9.7% to 18.2%; the fiscal deficit, the current account deficit and the external debt reached unsustainably high levels; and the economy entered into hyperinflation accompanied by a high incidence of capital flight. In 1985, the new government of Ángel Víctor Paz Estenssoro reacted to this situation with a radical policy change, introducing the “Nueva Política Económica” (New Economic Policy). First, a strict stabilization program, which included devaluing the local currency and implementing a restrictive monetary and fiscal policy, was carried out. Due to these policy measures, Bolivia quickly regained its internal and external macroeconomic equilibrium.[1]

Second, the Bolivian government initiated a comprehensive and far-reaching structural reform process, which aimed at restoring growth by enhancing the allocative efficiency and the international competitiveness of product and factor markets (UDAPE 2001). In 1985, prices of goods and services as well as interest rates were liberalized, transactions in foreign currency were re-introduced, and the exchange-rate market was unified. In the labor market, freedom of contract and free collective bargaining between employers and employees were re-established, and fringe benefits and dismissal protection of public-sector workers were reduced. In 1986, the economy was opened up to foreign trade. A complex tariff structure in which tariff rates varied from 0% to 150% was replaced by a uniform tariff of 20% on all imports,[2] and most non-tariff barriers were eliminated. Furthermore, a tax reform was implemented, which reduced the number of taxes, simplified the collection mechanism, and broadened the tax base. In 1987, banking legislation was overhauled and modern institutions for the regulation and supervision of banks were put in place. In the same year, the "Instituto Nacional de Exportaciones" (National Export Agency) was created in

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1 The stabilization program is not further taken up in this paper; see Sachs and Larraín (1998) and Antelo (2000) for a more comprehensive treatment of this issue.

2 In 1990, the uniform tariff rate was reduced to 10% and later the tariff rate for capital goods (but not for consumer goods) was reduced to 5%.

order to promote and diversify Bolivian exports. In 1990, the Investment Act strengthened the rights of foreign investors by giving them investment guarantees, by providing equal treatment for domestic and foreign investors, and by creating settlement mechanisms for cross-border commercial disputes. In 1992, "Zonas Francas" (Special Economic Areas) were created, which offered tax incentives for manufacturing activities and eliminated import tariffs for inputs. In the same year, the Privatization Act provided the regulatory framework for the privatization of public enterprises and the disposal of other state-owned assets. In 1994, the scope of privatization was expanded to state monopolies.[3]

Burki and Perry (1997) and Morley et al. (1999) developed a set of reform indices, which can be used to compare the structural reform progress in different Latin American countries. Figure 1 depicts the indices for financial reform, tax reform, labor market deregulation, trade liberalization, capital account liberalization, and privatization for Bolivia and neighboring Chile.[4] Except for the trade and the capital account liberalization measures, Bolivia's reform indices were significantly lower than those of Chile in 1985. This is not surprising since Chile started reforming its economy already in 1978. By 1995,

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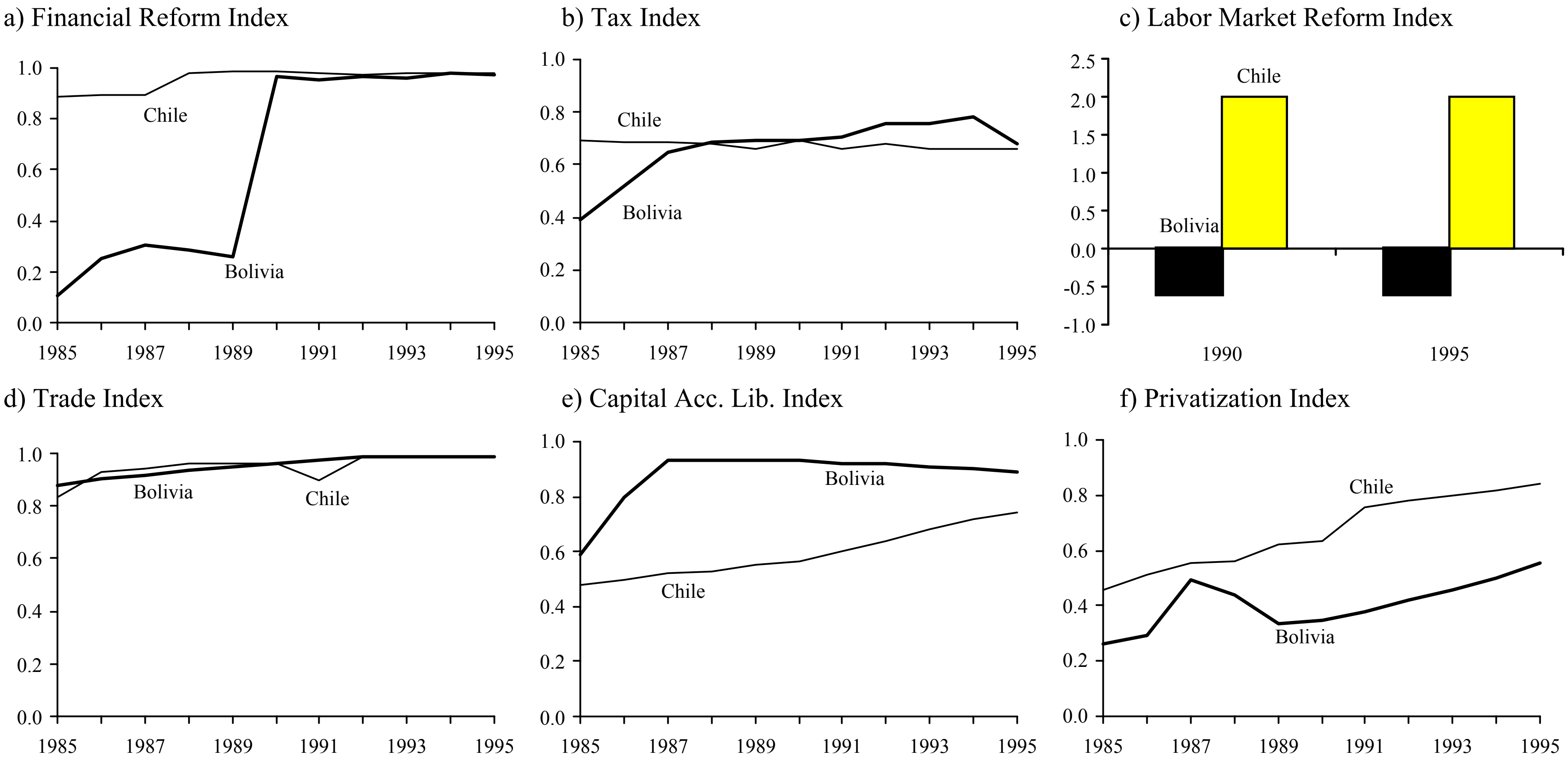
3 In the 1990s, Bolivia also implemented "second-generation" reforms, which aimed at improving the health and education system, at decentralizing the public administration, and at reforming the legal and judicial system. For a detailed description of these policy measures see UDAPE (2001).

4 Chile is used as benchmark country because it is often seen as the front-runner of successful structural reforms in Latin America.

however, Bolivia had caught up to or even overtaken Chile in all policy areas, except labor market reform and privatization. As concerns privatization, it has to be taken into account that Bolivia adopted a massive privatization program in the utilities sector in 1995/1996, which is not yet captured by the privatization index of the year 1995. It can, thus, be expected that this figure would be much higher for later years (Paunovic 2000). Against this background, Heinrigs and Steiner (2002) conclude that due to its fast structural reform process since 1985, Bolivia had achieved a reform level comparable to Chile by 2000. Only with respect to labor market reforms, Bolivia still lags substantially behind.

Figure 1: Structural Reforms in Bolivia: an Overview



Source: Authors’ calculations based on Burki and Perry (1997) and Morley et al. (1999).

2.2 Data

The data for the empirical part of our analysis comes from five household surveys collected by the Instituto Nacional de Estadísticas de Bolivia (National Statistical Office of Bolivia): the Encuestas Integradas de Hogares 1989, 1991, 1993 and 1995, and the Encuesta Nacional de Empleo 1997.[5] Household survey data are more appropriate than firm-level data for measuring earnings since the latter often do not include information on non-listed firms or micro-enterprises.

As labor income we define the reported wages and salaries of employees and the total earnings of self-employed and employers earned in their principal labor-market activity. This measure is deflated by the Consumer Price Index and divided by the reported working hours to obtain hourly wages at constant prices. Measurement problems may somewhat distort the reported labor incomes. First, fringe benefits could not be considered because the household surveys collect only the incidence and type, but not the monetary equivalent of fringe benefits. Second, total earnings of self-employed and employers may not always be measured net costs. The questionnaire does not contain enough detail to correct for this flaw. By contrast, income taxes hardly cause any significant distortions because they play only a negligible role in the revenues of the Bolivian

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5 As part of the MECOVI project, a joint program of World Bank, IADB and ECLAC for the improvement of surveys and the measurement of living conditions in Latin America and the Caribbean, the questionnaire of the Bolivian household surveys was comprehensively redesigned. As a result, the Encuesta Continua de Hogares 1999 could not be included in our analysis for lack of data compatibility.

government. As income-determining factors we use schooling years, age, gender, employment status (5 dummies), sectoral affiliation (12 dummies) and place of residence (9 dummies).

From all respondents we select those aged between 13 and 65 with strictly positive principal labor-market activity earnings and a full information set on the income-determining factors. We exclude those working as family workers and domestic servants because their reported labor incomes are especially prone to measurement errors. Since the Encuestas Integradas de Hogares (but not the Encuesta Nacional de Empleo) were only conducted in the departmental capitals and El Alto, the sample is restricted to urban areas of Bolivia.

3 Measuring Inequality

3.1 Methodology

We use two inequality measures: the Gini coefficient and the Atkinson index. Both measures are continuous, symmetric, mean independent (i.e., scale invariant) and satisfy Dalton's (1920) "principle of transfers".[6] In defining the inequality measures, $i \in [1, n]$ is the rank of the income unit when incomes are ordered from lowest to highest. y_i reflects the income and μ_y its empirical

6 The "principle of transfer" demands that a costless and rank-preserving transfer of income from a richer to a poorer income unit always result in a decrease in the inequality measure.

mean. ε is the degree of relative inequality aversion. The inequality measures are:

Gini coefficient: $G = \frac{-(n-1)}{n} + \frac{2}{n^2 \mu_y} \cdot \sum_{i=1}^{n} i\, y_i$, (1)

Atkinson index: $A(\varepsilon) = \begin{cases} 1 - \left[\frac{1}{n} \cdot \sum_{i=1}^{n} \left(\frac{y_i}{\mu_y} \right)^{1-\varepsilon} \right]^{\frac{1}{1-\varepsilon}} & if\, \varepsilon \neq 1 \\ 1 - \left(\prod_{i=1}^{n} \frac{y_i}{\mu_y} \right)^{\frac{1}{n}} & if\ \varepsilon = 1. \end{cases}$ (2)

The Gini coefficient represents two times the area between the 45° line and the Lorenz curve. Due to this straightforward graphical interpretation, the Gini coefficient has become the most commonly applied inequality measure. However, assuming a bell-shaped distribution, the Gini coefficient is insensitive to transfers within the group of low-ranking or within the group of high-ranking income units. This is because the sensitivity of the Gini coefficient to transfers between two income units depends only on their ranks but not on their income levels.

The Atkinson index has three desirable features. First, it can be derived axiomatically to be consistent with a social welfare maximization model. Second, by varying a single parameter, the Atkinson index encompasses an entire family of social welfare functions ranging from completely egalitarian to very non-egalitarian ones. The higher the degree of relative inequality aversion,

ε, the more weight the index attaches to the lower tail of the distribution and the more sensitive it becomes to low incomes. Following Atkinson (1970), we assume that ε lies within the range (0, 2.5). Third, the Atkinson index has an intuitive monetary interpretation. If incomes were equally distributed, the equally distributed equivalent income

$$\tilde{y} = [1 - A(\varepsilon)] \cdot \mu_y, \tag{3}$$

would provide the same level of social welfare as the actual income distribution.

3.2 Empirical Results

Table 1 shows that three out of four inequality indices, namely *G*, $A(0.5)$, and $A(1.5)$, increased till 1995 and declined slightly thereafter, while $A(2.5)$ followed a more irregular time pattern. Focusing on the 1989-to-1997 percentage growth rates (denoted by ^), we observe that all inequality indices increased over the observation period, indicating a rise in wage inequality. The percentage growth rate of wage inequality was highest for $A(0.5)$ and lowest for $A(2.5)$. Since the latter Atkinson index attaches more weight to low-incomes, this result suggests that the rise in wage inequality was most pronounced at the upper tail of the distribution. In other words, changes in income distribution mainly occurred from middle to high incomes, rather than from low to middle incomes.

Table 1: Income Inequality in Urban Bolivia 1989–1997

	1989	1991	1993	1995	1997		1989-1997
G	0.497	0.518	0.528	0.543	0.532	$\hat{G}$	7.18%
A(0.5)	0.210	0.228	0.232	0.247	0.235	$A(\hat{0}.5)$	12.12%
A(1.5)	0.471	0.500	0.514	0.523	0.521	$A(\hat{1}.5)$	10.54%
A(2.5)	0.677	0.743	0.699	0.690	0.711	$A(\hat{2}.5)$	5.04%

Source: Authors' calculations.

4 Decomposing Income Inequality

4.1 Methodology

After having measured income inequality, we proceed by "decomposing" it using the Fields' (2001) decomposition methodology. There are two questions to be answered: (a) the "level question": What fraction of income inequality is accounted for by each explanatory variable? (b) the "difference question": What fraction of the change in income inequality between one date and another can these variables explain?

We start by addressing the "level question". In a standard income-generating function

$$y_t = a_t{}'Z_t = [\alpha_t \ \beta_{1t} \ \dots \ \beta_{jt} \ \dots \ \beta_{Jt} \ 1]' \cdot [1 \ x_{1t} \ \dots \ x_{jt} \ \dots \ x_{Jt} \ \varepsilon_t], \tag{4}$$

we define y_t as a vector of log incomes, and Z_t as a matrix of the constant, J explanatory variables and the error term. Given equation (4), it can be shown that for any continuous and symmetric inequality measure $I(y_{1t},....,y_{nt})$ which is zero for equally distributed incomes, the absolute inequality-level weight

$$s_{jt} := \frac{\operatorname{cov}\left(a_{jt} Z_{jt}, y_t\right)}{\sigma_{yt}^2} = \frac{a_{jt} \cdot \sigma_{Z_j t} \cdot \operatorname{cor}\left(Z_{jt}, y_t\right)}{\sigma_{yt}}, \tag{5}$$

where $\sum_{j=0}^{J} s_{jt} = R_t^2$ and $\sum_{j=0}^{J+1} s_{jt} = 1$,

is the fraction of total income inequality that is attributable to the j^{th} explanatory variable.[7]

Next, we turn to the "difference question". The fraction of the change in a particular inequality measure over time period *dt* which is explained by the j^{th} income-determining factor is given by its inequality-change weight

$$\pi_{jdt}(I) := \frac{s_{jt_0+dt}\, I_{t_0+dt} - s_{jt_0}\, I_{t_0}}{I_{t_0+dt} - I_{t_0}}, \tag{6}$$

where $\sum_{j=1}^{J} \pi_{jdt}(I) = 1$.

Finally, the inequality-change weights can be further decomposed into three components: the fractions attributable to (a) changes in the regression coefficients (hereafter referred to as *coefficient effect*), (b) changes in the correlation coefficient between the explanatory variable and log incomes (*correlation effect*), and (c) changes in the variance of the regressor relative to the variance of log incomes (*variance effect*). Logarithmically differentiating (5) and dividing both sides of the equation by $\hat{s}_{jt}$ yields

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[7] Upon dividing by the coefficient of determination R_t^2, we obtain the *relative* inequality-level weights $p_{jt} := \frac{s_{jt}}{R_t^2}$, where $\sum_{j=0}^{J} p_{jt} = 1$, i.e., the fraction of the *explained* income inequality that is attributable to the j^{th} explanatory variable.

$$\underset{a)}{\frac{\hat{a}_{jt}}{\hat{s}_{jt}}} + \underset{b)}{\frac{\hat{\text{cor}}\left[Z_{jt}, Y_t\right]}{\hat{s}_{jt}}} + \underset{c)}{\frac{\widehat{\left(\sigma_{Z_{jt}} / \sigma_{yt}\right)}}{\hat{s}_{jt}}} = 1, \tag{7}$$

with ^ again indicating percentage growth rates. Equation (7) is intuitively appealing. The j^{th} regressor in the income-generating function contributes more to explaining an observed increase in inequality (a) the larger the increase in its regression coefficient, (b) the larger the increase in its correlation coefficient with income and (c) the larger the increase in the variance of the regressor relative to the variance of log incomes.

4.2 Empirical Results

The input for the Fields decomposition was obtained by running a regression of the log hourly wages on a constant, age, age squared, schooling years and dummies for gender, employment status, sectoral affiliation and place of residence:

$$\begin{aligned} \ln(yph_t) = \alpha \\ &+ \beta_1 \text{age}_t + \beta_2 \text{age}^2{}_t + \beta_3 \text{school}_t + \\ &+ \beta_4 \text{d\_sex}_t + \sum_{q=1}^{5} \beta_{q+4} \text{d\_ocu}_{qt} + \sum_{r=1}^{11} \beta_{r+9} \text{d\_sec}_{rt} + \sum_{s=1}^{8} \beta_{s+20} \text{d\_dep}_{st} \\ &+ \varepsilon_t . \end{aligned} \tag{8}$$

We estimate equation (8) by weighted least squares using White's (1980) heteroskedasticity-consistent covariance matrix estimator to calculate standard errors.[8] Based on the regression results, we carry out Fields' (2001) "level decomposition" calculating the absolute inequality-level weights for each of the explanatory variables. The results are summarized in Table 2.[9]

Table 2: Inequality-level Weights s_{jt} (%)

	1989	1991	1993	1995	1997
Age	4.60	5.30	5.01	4.06	4.19
Education	11.42	12.69	22.60	16.22	19.61
Gender	1.17	1.72	1.68	1.03	1.05
Employment Status	2.63	2.14	5.58	3.54	4.07
Sectoral Affiliation	3.61	3.47	4.77	3.83	4.39
Residence	4.23	2.17	3.39	3.59	2.62
Sum 1 = R_t^2	27.66	27.49	43.03	32.27	35.93
Residual	72.35	72.51	56.97	67.73	64.07
Sum 2 = 100	100.01	100.00	100.00	100.00	100.00

Source: Authors' calculations.

8 The regression results are presented in detail in Appendix A.

9 For the case of Bolivia, the "level question" but not the "difference question" of Fields' (2001) decomposition methodology was also addressed by Fields et al. (1998) for the years 1992–1995, and by Andersen (1999) for the years 1989–1995.

The coefficient of determination of the regression model, R_t^2, which is identical to the sum of the inequality-level weights of all explanatory variables rose from 27.7% in 1989 to 35.9% in 1997. In other words, the fit of the regression model improved and the explanatory power of the income-determining factors increased substantially in the 1990s. *Education* was by far the most important income-determining factor in every single observation period. Over time, this variable could further "extend the lead". Its inequality-level weight rose from 11.4% in 1989 to 19.6% in 1997. The second most important explanatory variable was *Age.*[10] Its explanatory power varied between 4.1% and 5.3% with no discernible trend over time. The *Residence* dummies also played a significant, though declining role in explaining wage inequality. This income-determining factor accounted for 4.2% in 1989, but only 2.6% in 1997. The *Employment Status* and the *Sectoral Affiliation* dummies both gained importance. Their inequality-level weights rose from 2.6% to 4.1% and from 3.6% to 4.4%, respectively. The role of the *Gender* dummy was negligible, accounting for only between 1.0% and 1.7% of total wage inequality.

In the second step of the analysis, we calculate the contribution of each explanatory variable to the changes in the inequality indices according to equation (6). In Table 3, we compile the 1989-to-1997 inequality-change weights $\pi_{jdt}(I)$ for the Gini coefficient and the Atkinson indices.

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10 Under *Age* we subsumed the regressors age and age squared in equation (8).

Table 3: 1989-to-1997 Inequality-Change Weights $\pi_{jdt}(I)$ (%)

	G	*A(0.5)*	*A(1.5)*	*A(2.5)*
Age	-1.53	0.80	0.29	-3.95
Education	133.80	87.28	97.40	182.32
Gender	-0.59	0.08	-0,06	-1.28
Employment Status	24.22	16.01	17.80	32.79
Sectoral Affiliation	15.24	10.82	11.78	19.85
Residence	-19.78	-10.65	-12.64	-29.30
Residual	-51.36	-4.34	-14.57	-100.42
Sum = 100	100.00	100.00	100.00	100,01

Source: Authors' calculations.

Fields' (2001) "difference decomposition" shows that *Education* was the main contributor to the 1989-to-1997 rise in wage inequality. Depending on the inequality index used, this variable alone accounted for between 87.3% and 182.3% of the empirically observed rise in the inequality measures. The fact that its inequality-change weight for $A(2.5)$ was more than twice as large as the one for $A(0.5)$ reveals that the influence of *Education* on wages rose particularly at the lower tail of the distribution. Other variables that contributed to the rise in wage inequality are the *Employment Status* and the *Sectoral Affiliation* dummies. However, even taken together, their explanatory power was small compared to *Education*. Between 1989 and 1997, *Age* and the *Gender* dummy had no influence on the change in wage inequality once other income-determining factors are controlled for. Their inequality-change weights were close to zero.

The rise in wage inequality was counteracted by the *Residence* dummies. Other things being equal, this income-determining factor would have reduced the inequality measures by between 10.7% and 29.3% of the empirically observed 1989-to-1997 rise in wage inequality. However, the main factor moderating the rise in wage inequality was the *Residual*. Depending on the inequality index used, its inequality-change weights were between –4.3% and –100.4%. This influence is due to the improvement of the fit of the regression model from 1989 to 1997. The role of unobserved variables in wage setting declined, thereby reducing wage inequality. Again we observe that this effect was most pronounced at the lower tail of the distribution.

In order to understand the transmission mechanisms behind the dominance of *Education* in explaining the rise in wage inequality, we decomposed the inequality-change weight of this variable for the period 1989–1997 according to equation (7). Table 4 shows that the increase in the correlation coefficient between *Education* and wages (correlation effect) accounted for 60.0% of *Education*'s contribution to the rise in Bolivian wage inequality. The increase in the inequality of schooling years (variance effect) explained another 27.5%. In contrast, the component weight of the coefficient effect, which captures the rise in the return to schooling, was only 14.4%.[11]

[11] This result is at odds with comparable studies of the United States, where the coefficient effect clearly dominates the other two effects (Fields 2001).

Table 4: *Education*'s Component Weights (%)[12]

a) Coefficient Effect	14.43
b) Correlation Effect	59.97
c) Variance Effect	27.46
Sum = 100%	101.86

Source: Authors' calculations.

5 Explaining Income Inequality

5.1 Stylized Facts

The empirical results on recent trends in Bolivian wage inequality can be summarized in three stylized facts.

- Between 1989 and 1997, there was a rise in wage inequality which was most pronounced at the upper tail of the distribution.
- *Education* gained and omitted variables lost influence in wage setting. As a result, *Education* was the main contributor and omitted variables were the main counteractor to the rise in wage inequality. These effects were especially strong at the lower tail of the distribution.

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12 Due to two approximation errors, the three component weights do not add up to 100%. First, real-world changes in each component of equation (7) are non-infinitesimal. Second, a_{jt} and $\mathrm{cor}(Z_{jt}, y_t)$ are both functions of $\mathrm{cov}(Z_{jt}, y_t)$ so that one component can not be varied without the other (Fields 2001).

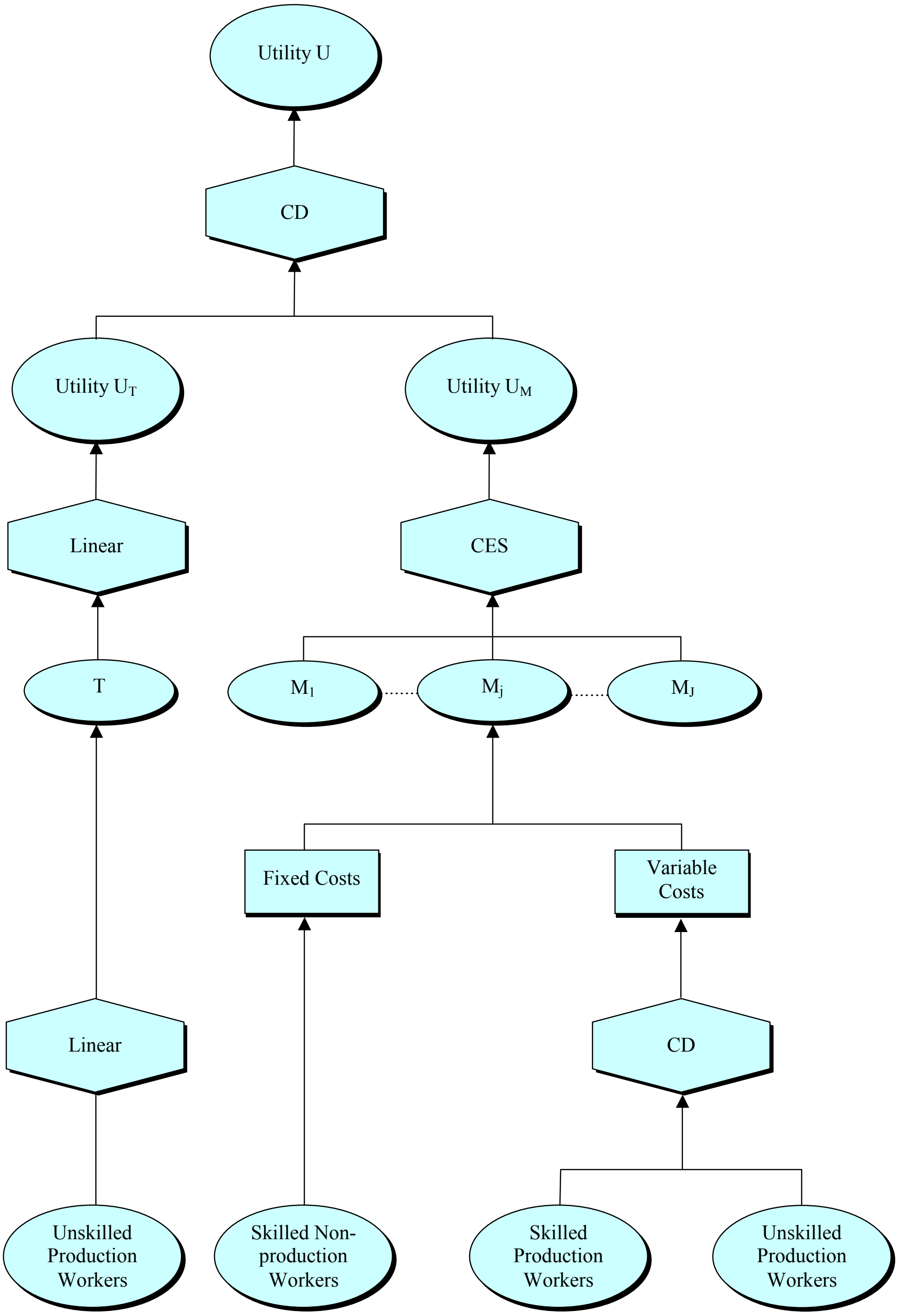
- A large share of *Education*'s contribution to the rise in wage inequality can be attributed to a rise in the correlation coefficient between wages and schooling years (correlation effect).

In order to understand the driving forces and the transmission mechanisms behind the empirical results, we set up a simple general equilibrium model. We show that market imperfections enable workers who are covered by labor market institutions to appropriate rents. Structural reforms are fundamentally about reducing market imperfections. Hence, they erode the favorable income position of these workers, thereby, changing the wage and employment distribution of the whole economy. We argue that labor market institutions are biased towards the middle class. We incorporate this notion into our model by making two assumptions. (a) labor market institutions are not present in the informal sector of the economy. (b) in the formal sector, they only cover unskilled workers. In order to be compatible to the standard new-keynesian terminology, we use the term "union" to refer to all those labor market institutions that give rise to bargaining power in wage negotiations.

5.2 The Basic Model

As shown in Figure 2, we assume a dual economy. The market structure of the informal sector, which produces the traditional good *T*, is perfectly competitive. In the formal sector, *J* monopolistically competitive firms produce *J* varieties of the modern good *M*.

Figure 2: Utility and Production Tree



On the first stage of the utility maximization problem, household $i = 1,\dots,N$ solves

$$
\begin{aligned}
&\max_{\{M_i,T_i\}} U_i = M_i^{\mu_i} T_i^{1-\mu_i}, \quad 0 < \mu_i < 1, \\
&\text{s.t.} \quad Y_i = P_M M_i + P_T T_i, \\
&\text{with} \quad Y_i = \begin{cases} W_{M_j}, \text{if } i \text{ is employed by firm } j \text{ in the formal sector} \\ W_T, \text{if } i \text{ is employed in the informal sector} \end{cases}
\end{aligned}
\tag{9}
$$

allocating income Y_i to the consumption of M and T according to

$$
\frac{M_i}{T_i} = \frac{\mu_i}{1-\mu_i} \cdot \frac{P_T}{P_M}. \tag{10}
$$

The household loves variety in the modern sector and derives utility from J varieties of the modern good according to the CES utility function

$$
M_i = \left[J^{-\frac{1}{\eta}} \sum_{j=1}^{J} M_{ij}^{\frac{\eta-1}{\eta}} \right]^{\frac{\eta}{\eta-1}}, \quad M_{ij} \geq 0 \;\forall\, i, j, \quad \eta > 1, \tag{11}
$$

where η is the absolute value of the elasticity of substitution between the different varieties of the modern good M and reflects the degree of market competition in the formal sector. Maximizing utility subject to the budget constraint and aggregating over all households yields the demand function of variety j

$$M_j = \left(\frac{P_{M_j}}{P_M} \right)^{-\eta} \cdot \frac{Y_M}{P_M J}, \tag{12}$$

where Y_M is the income share spent on the modern good, P_{M_j} is the price of variety *j,* and

$$P_M = \left[\frac{1}{J} \sum_{j=1}^{J} P_{M_j}^{1-\eta} \right]^{\frac{1}{1-\eta}} \tag{13}$$

is the price index for the modern good. The price elasticity of demand is given by

$$\varepsilon_{M_j, P_{M_j}} = -\eta \,. \tag{14}$$

The output of variety *j* depends on the number of unskilled production workers, L_{M_j}, and the number of skilled production workers, $H^v_{M_j}$, and is produced with a Cobb-Douglas production function

$$M_j = L_{M_j}{}^{\alpha} \cdot H^v_{M_j}{}^{1-\alpha} \,. \tag{15}$$

Additionally, the workforce of firm *j* consists of skilled non-production workers, $H^f_{M_j}$, who are assumed to receive the same wage as the skilled production workers.

Maximizing firm *j's* profits[13] subject to its demand function (12) implies that the wages of unskilled workers, W_{M_j}, and the wages of skilled workers, Q_{M_j}, are set according to

$$W_{M_j} = \frac{\eta - 1}{\eta} \cdot \alpha \cdot \left(\frac{H^v_{M_j}}{L_{M_j}} \right)^{1-\alpha} \cdot P_{M_j} \text{, and} \tag{16}$$

$$Q_{M_j} = \frac{\eta - 1}{\eta} \cdot (1-\alpha) \cdot \left(\frac{H^v_{M_j}}{L_{M_j}} \right)^{-\alpha} \cdot P_{M_j} \, . \tag{17}$$

The wage elasticity of unskilled labor demand is given by

$$\varepsilon_{L_{M_j}, W_{M_j}} = -1 - \alpha(\eta - 1) \; . \tag{18}$$

In the informal sector, labor productivity of workforce L_T is constant and normalized to one, i.e.,

$$T = L_T \, . \tag{19}$$

In perfectly competitive product markets, firms set prices equal to marginal costs, which implies

$$P_T = W_T \, , \tag{20}$$

where W_T is the wage paid in the informal sector.

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13 The profits of firm *j* are $\Pi_{M_j} = P_{M_j} M_j - W_{M_j} L_{M_j} - Q_j \left(H^v_{M_j} + H^f_{M_j} \right)$.

The economy is populated with risk-neutral individuals who supply labor inelastically. Unskilled labor is employed in both sectors. In the informal sector, the unskilled labor market is atomistic, whereas in the formal sector, unskilled wages result from negotiations between unions and firms.[14] Skilled labor is employed in the formal sector only. Due to the lack of an outside option, skilled workers are also assumed to be wage takers.

The wage bargain takes place in a right-to-manage set-up,[15] where each firm negotiates with a single in-house union[16] at the beginning of each period (see Figure 3).

The negotiation partners' stake in the wage bargaining is the difference in payoffs between a situation with and without an agreement. Union *j* is assumed to represent only firm *j's* unskilled workers. Upon successful completion of the negotiations, union *j* gains a rent of

$$\Gamma_{U_j} = L_{M_j} \cdot \left(W_{M_j} - Z\right), \tag{21}$$

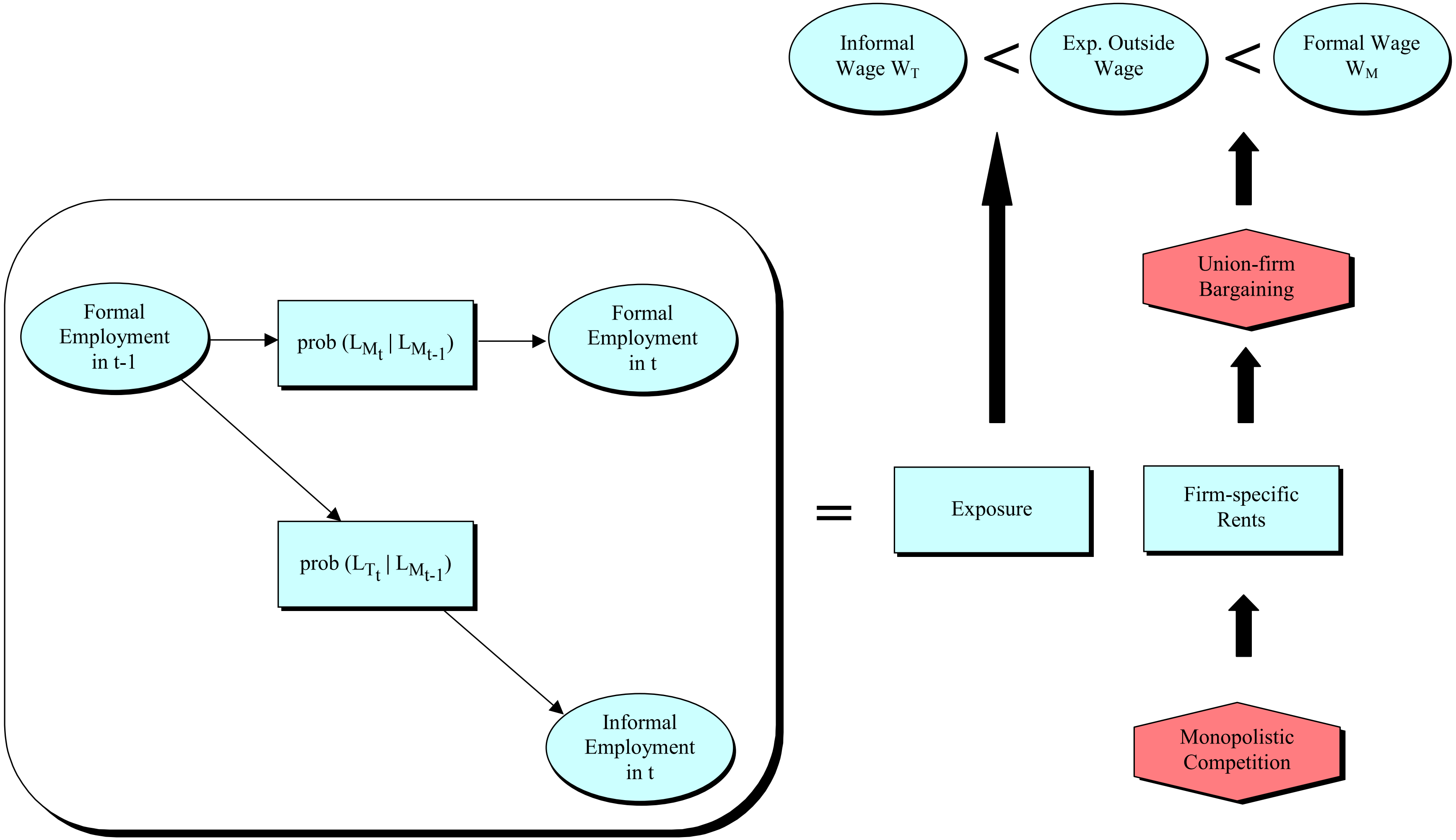
where Z is the expected outside wage.

[14] This assumption can be justified on the grounds that (a) rents only accrue in the monopolistically competitive formal sector and (b) unskilled formal-sector workers can appropriate a share of these rents since demand for unskilled labor is decreasing in wages.

[15] In other words, the two negotiation partners jointly determine the wage, while the firm unilaterally sets the employment level afterwards.

[16] As a reminder, if we speak of unions, we mean all those labor-market institutions that give rise to bargaining power in wage negotiations.

Figure 3: The Wage Setting Tree for Unskilled Formal-Sector Workers



Due to its monopoly power, firm *j* can set prices as a mark-up on marginal costs. Its stake in the wage bargaining is equal to its variable profits

$$\Gamma_{M_j} = P_{Mj} \cdot L_{M_j}{}^{\alpha} \cdot H_{M_j}^{v}{}^{1-\alpha} - W_{M_j} \cdot L_{M_j} - Q_{M_j} \cdot H_{M_j}^{v} . \tag{22}$$

Assuming an asymmetric Nash bargaining solution, the wage is set to maximize the geometric average of the negotiation partners' rents from reaching an agreement

$$\Omega_j = \Gamma_{U_j}^{\beta_j} \, \Gamma_{M_j}^{1-\beta_j} , \tag{23}$$

where β_j is unskilled workers' share of the Nash Maximand Ω_j and reflects the bargaining power of union *j*. Since firms can choose employment ex-post, the negotiation partners maximize the Nash Maximand by choosing the wage equal to[17]

$$W_{M_j} = \left[1 + \frac{\beta_j}{\alpha \cdot (\eta - 1)} \right] \cdot Z . \tag{24}$$

The wage is set as a mark-up on the expected outside wage.

We assume that after the wage bargaining is completed at the end of period *t-1*, all unskilled jobs of the formal sector are newly allocated. Members of union *j* who are not re-employed by firm *j* expect either to find employment in one of the other *J-1* formal-sector firms at the average formal-sector wage W_M

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17 See Appendix B1.

or to have to work in the informal sector at W_T. The expected outside wage in period *t* is, thus, given by

$$Z = \left(1 - prob(L_{T_t} \middle| L_{M_{t-1}})\right) \cdot W_M + prob(L_{T_t} \middle| L_{M_{t-1}}) \cdot W_T, \tag{25}$$

where $prob(L_{T_t} | L_{M_{t-1}})$ is the conditional probability that an unskilled worker who was employed in the formal sector in period *t-1* (hereafter referred to as *ex-formal-sector unskilled worker*) has to work in the informal sector in period *t*. Assuming further that all unskilled workers are equally likely to find employment in the formal sector, equation (25) simplifies to[18]

$$Z = \frac{L_M}{L} \cdot W_M + \left(1 - \frac{L_M}{L}\right) \cdot W_T. \tag{26}$$

We consider a symmetric equilibrium in which all households have identical preferences, and in which union bargaining power and the number of non-production workers are equal for all firms. There are no administrative barriers to market entry[19] and the economy is assumed to be closed.

Using these equilibrium conditions and collecting terms, we arrive at three equations to simulate the model[20]

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18 This assumption will be relaxed in Section 5.3..

19 Hence, the rents earned by formal-sector firms just cover the wages of their high-skilled non-production workers so that firms' profits are zero, too.

20 See Appendix B2.

$$\omega := \frac{W_M}{Q_M} = \frac{\alpha \cdot (\eta - 1)}{\eta - \alpha \cdot (\eta - 1)} \cdot \frac{H}{L} \cdot \left(\frac{L_M}{L} \right)^{-1}, \tag{27}$$

$$1 + \theta := \frac{W_M}{W_T} = \frac{\mu}{1 - \mu} \cdot \frac{\eta - 1}{\eta} \cdot \alpha \cdot \frac{1 - \frac{L_M}{L}}{\frac{L_M}{L}}, \tag{28}$$

$$1 + \theta := \frac{W_M}{W_T} = \frac{\left[1 + \frac{\beta}{\alpha \cdot (\eta - 1)} \right] \cdot \left(1 - \frac{L_M}{L} \right)}{1 - \left[1 + \frac{\beta}{\alpha \cdot (\eta - 1)} \right] \cdot \frac{L_M}{L}}, \tag{29}$$

where ω represents the relative wage of unskilled workers in the formal sector, θ is the sectoral wage premium for unskilled formal-sector workers, and $\frac{L_M}{L}$ is the unskilled employment share of the formal sector. Equation (27) balances the costs and benefits of substituting skilled for unskilled labor in the formal sector. Equations (28) and (29) represent the quasi-demand and quasi-supply function for unskilled labor in the formal sector.

The labor market equilibrium depends on five model parameters: the relative skill endowment, $\frac{H}{L}$, the income share spent on the modern good, μ, the degree of competition in the formal sector, η, the share parameter of the production function in the formal sector, α, and the union bargaining power, β.

Figure 4 depicts the comparative statics of the model with $\frac{H}{L} = 0.15$, $\mu = 0.4$, $\eta = 3$, $\alpha = 0.5$, and $\beta = 0.15$ as baseline parameters.[21]

As was to be expected, an increase in the relative skill endowment, $\frac{H}{L}$, causes an increase in the relative wage of unskilled workers in the formal sector ($\omega \uparrow$). Given the Cobb-Douglas production function in the formal sector, the relationship between the two variables is linear, and the sectoral wage premium for unskilled formal-sector workers and the unskilled employment share of the formal sector are independent of the relative skill endowment.

A rise in the income share spent on the modern good, μ, increases both the price and the quantity demanded of the modern good. Both effects raise the rents earned in the formal sector. The second effect also causes an increase in the quantities demanded and the wages of skilled and unskilled formal-sector workers. In the case of unskilled workers, the rise in wages is partly offset by a movement of workers from the informal to the formal sector ($\frac{L_M}{L} \uparrow$). Consequently, the relative wages of unskilled workers in the formal sector

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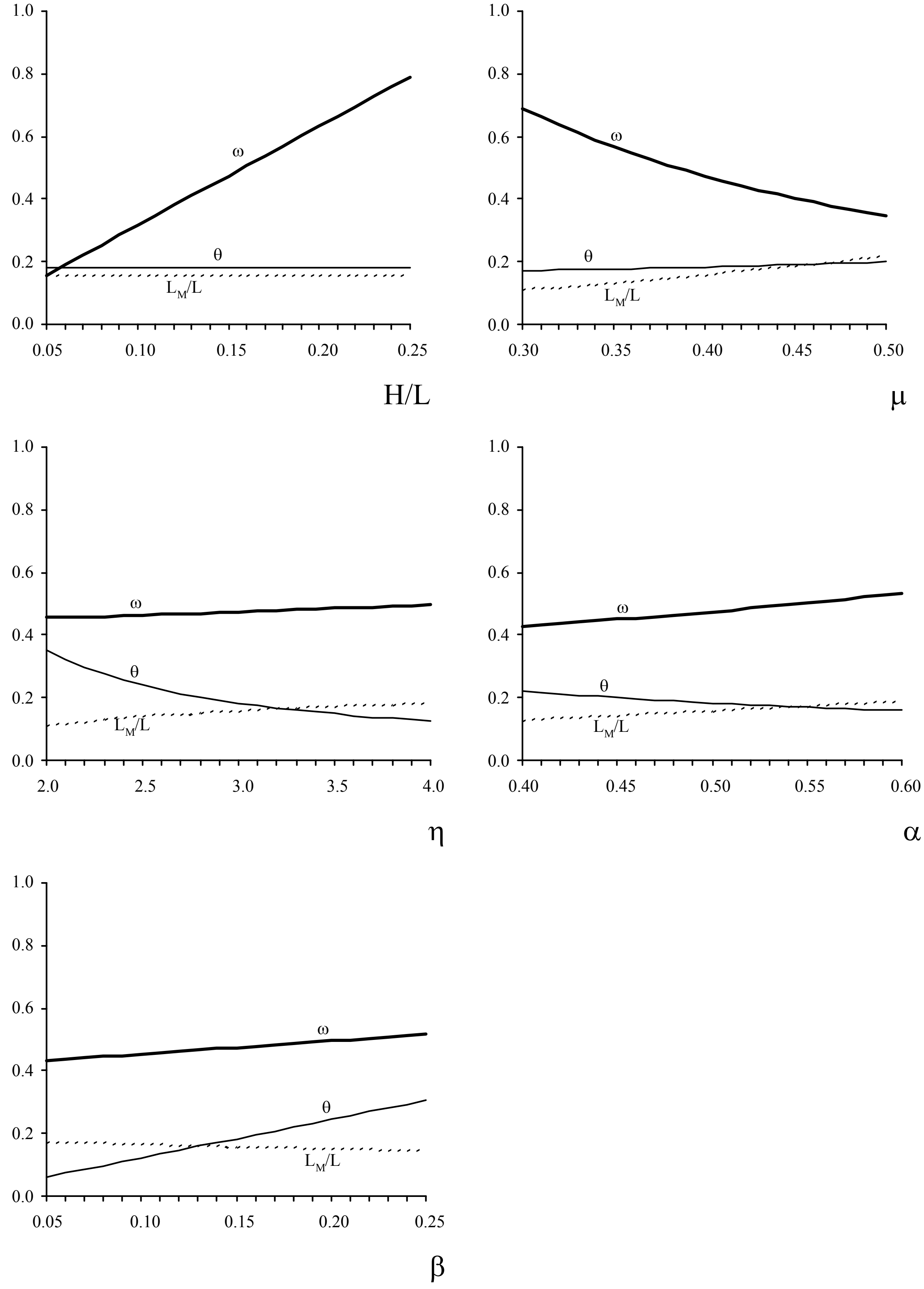
21 The model puts few restrictions on the parameters and is well-behaved to parameter changes so that the model results are not sensitive to the choice of the baseline parameters.

Figure 4: Simulation Results of the Basic Model



Source: Authors' calculations.

declines ($\omega \downarrow$). For a unitary elasticity of substitution between skilled and unskilled workers in the formal sector, the rise in $\frac{L_M}{L}$ is smaller than the rise in total rents earned in the formal sector. Hence, the sectoral wage premium of unskilled formal-sector workers increases ($\theta \uparrow$).

An increase in the degree of competition in the formal sector, η, reduces the price and raises the quantity demanded of the modern good. For both reasons, the rents earned in this sector decrease. The sectoral wage premium of unskilled formal-sector workers declines ($\theta \downarrow$). At the same time, the number of firms and, thus, of skilled non-production workers in the formal sector falls, thereby, increasing the supply of skilled production workers. The increase in the quantity demanded of the modern good additionally raises the demand for skilled and unskilled formal-sector workers. The sum of these partial effects causes the employment share of the formal sector to expand ($\frac{L_M}{L} \uparrow$) and the relative wage of unskilled workers in the formal sector to rise ($\omega \uparrow$).

A rise in the share parameter of the production function in the formal sector, α, reduces the productivity gap between skilled and unskilled workers. Consequently, demand for unskilled formal-sector workers rises resulting in an increase of ω and $\frac{L_M}{L}$. The increase of the unskilled employment share of the formal sector reduces the rent per unskilled formal-sector worker and, thus, their sectoral wage premium ($\theta \downarrow$).

Finally, an increase in union bargaining power increases the proportion of rents going to unskilled formal-sector workers. As a result, the wage gap to skilled formal-sector workers narrows ($\omega\uparrow$) and the wage gap to unskilled informal-sector workers widens ($\theta\uparrow$). Since unskilled formal-sector workers become more expensive, they are replaced by skilled workers. Hence, the unskilled employment share of the formal sector goes down ($\frac{L_M}{L}\downarrow$).

The comparative statics of the model can be used to discuss the impact of structural reforms on wages and employment and, thus, on wage inequality. However, before we proceed towards this goal, we develop two more elaborate versions of the basic model that capture two important features of wage setting.

5.3 Two More Elaborate Models

Saint-Paul (2000) pointed out that the size and distribution of rents is determined by the degree of substitution between production factors. To incorporate this idea into our model, we replace equation (15) by the CES production function

$$M_j = \left[\alpha \cdot L_{M_j}^{\frac{\sigma-1}{\sigma}} + (1-\alpha) \cdot H_{M_j}^{\nu\,\frac{\sigma-1}{\sigma}}\right]^{\frac{\sigma}{\sigma-1}}, \tag{30}$$

where σ is the elasticity of substitution between skilled and unskilled workers in the formal sector. Using this model specification, the wage elasticity of unskilled labor in the formal sector reads

$$\varepsilon_{L_{M_j},W_{M_j}} = -\eta + (\eta - \sigma) \cdot \frac{\frac{1-\alpha}{\alpha} \cdot \left(\frac{H_{M_j}^{v}}{L_{M_j}}\right)^{\frac{\sigma-1}{\sigma}}}{\left[1 + \frac{1-\alpha}{\alpha} \cdot \left(\frac{H_{M_j}^{v}}{L_{M_j}}\right)^{\frac{\sigma-1}{\sigma}}\right]}, \tag{31}$$

and we arrive at the following system of equations to jointly determine the three endogenous variables of the model

$$\omega = \left[\frac{(\eta - 1) \cdot \omega^{\sigma-1}}{1 + \eta \cdot \left(\frac{1-\alpha}{\alpha}\right)^{\sigma} \cdot \omega^{\sigma-1}}\right]^{\frac{1}{\sigma}} \cdot \left(\frac{H}{L}\right)^{\frac{1}{\sigma}} \cdot \left(\frac{L_M}{L}\right)^{-\frac{1}{\sigma}}, \tag{32}$$

$$1 + \theta = \frac{\mu}{1-\mu} \cdot \frac{\eta - 1}{\eta} \cdot \left[1 + \left(\frac{1-\alpha}{\alpha}\right)^{\sigma} \cdot \omega^{\sigma-1}\right]^{-1} \cdot \frac{1 - \frac{L_M}{L}}{\frac{L_M}{L}}, \tag{33}$$

$$1 + \theta = \frac{\left[1 + \frac{(\eta - 1)}{\beta} + \sigma \cdot \left(\frac{1-\alpha}{\alpha}\right)^{\sigma} \cdot \omega^{\sigma-1}\right] \cdot \left(1 - \frac{L_M}{L}\right)}{\frac{(\eta - 1)}{\beta} + \left(\frac{1-\alpha}{\alpha}\right)^{\sigma} \cdot \omega^{\sigma-1} - \left[1 + \frac{(\eta - 1)}{\beta} + \sigma \cdot \left(\frac{1-\alpha}{\alpha}\right)^{\sigma} \cdot \omega^{\sigma-1}\right] \cdot \frac{L_M}{L}}. \tag{34}$$

The impact of this change in the model specification on the previous results can be summarized as follows. First, the influence of $\frac{H}{L}$ on θ and $\frac{L_M}{L}$ is no longer zero. For $\sigma < 1$, θ is decreasing, and $\frac{L_M}{L}$ is increasing in $\frac{H}{L}$, while the reverse holds true for $\sigma > 1$.[22] Second, a rise in μ can reduce θ for $\sigma >> 1$. And third, ω can be decreasing in η for $\sigma << 1$. Hence, in contrast to the basic model, the effects of $\frac{H}{L}$ on θ and $\frac{L_M}{L}$, of μ on θ, and of η on ω are now undetermined.

Additionally, we analyze the impact of changes in the elasticity of substitution between skilled and unskilled workers in the formal sector on ω, θ, and $\frac{L_M}{L}$. A low value of σ means that skilled and unskilled workers interact closely in the production process of the formal sector and that relatively scarce factors are highly productive. Hence, the demand for unskilled workers is low, which implies that both the relative wage of unskilled workers and the unskilled employment share of the formal sector are low, too. Being few in numbers, unskilled formal-sector workers can appropriate a high rent per head and, thus, receive a high sectoral wage premium. As a result, as shown in Figure 5, an increase in σ raises ω and $\frac{L_M}{L}$, but reduces θ.

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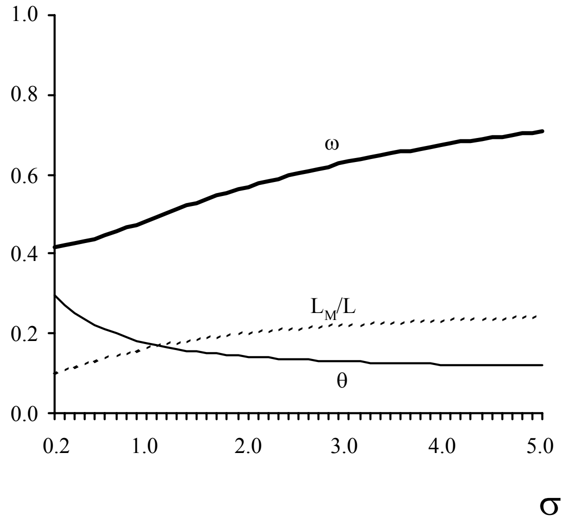
22 Additionally, the relationship between ω and $\frac{H}{L}$ is no longer linear, but convex for $\sigma < 1$, and concave for $\sigma > 1$.

Figure 5: Simulation Results of the Elaborate Model 1



Source: Authors' calculations.

In the basic model, we made the strong assumption that regardless of their sectoral affiliation in period *t-1*, all unskilled workers are equally likely to find formal-sector employment at the beginning of period *t*. Due to sector-specific human capital and employment protection, however, it is more reasonable to argue that formal-sector workers are relatively little exposed to the risk of having to work in the informal sector. To account for differences in relative exposure, we assume that at the beginning of period *t*, the probability of having to work in the informal sector is smaller for ex-formal-sector workers (= insiders) than for ex-informal-sector workers (= outsiders), i.e.,

$$prob(L_{T_t} \mid L_{M_{t-1}}) = \psi \cdot prob(L_{T_t} \mid L_{T_{t-1}}), \tag{35}$$

where $0 < \psi < 1$ measures the degree of relative exposure. In this model specification, the expected outside wage is given by

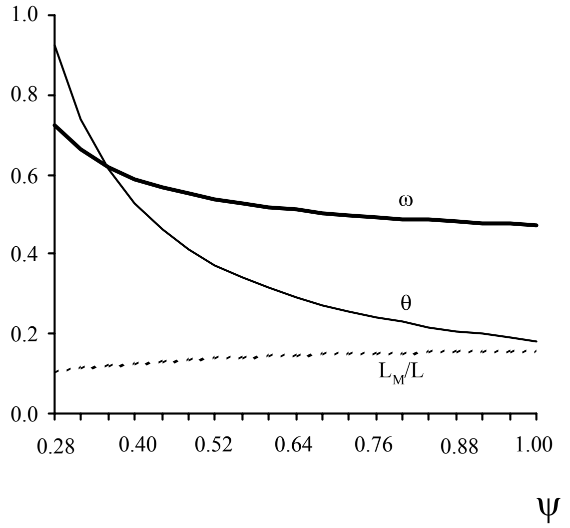
$$Z=\left(1-\frac{1-\frac{L_M}{L}}{\frac{L_M}{L}+\left(1-\frac{L_M}{L}\right)\cdot\frac{1}{\psi}}\right)\cdot W_V+\left(\frac{1-\frac{L_M}{L}}{\frac{L_M}{L}+\left(1-\frac{L_M}{L}\right)\cdot\frac{1}{\psi}}\right)\cdot W_T\,. \qquad (36)$$

Returning to a Cobb-Douglas production function in the formal sector, equations (26) and (27) from the basic model still hold and the new quasi-supply function for unskilled labor in the formal sector reads

$$1+\theta:=\frac{W_M}{W_T}=\frac{\left[1+\frac{\beta}{\alpha\cdot(\eta-1)}\right]\cdot\left(1-\frac{L_M}{L}\right)}{\left(1-\frac{L_M}{L}\right)-\frac{\beta}{\alpha\cdot(\eta-1)}\cdot\left[\frac{L_M}{L}+\left(\frac{1-\psi}{\psi}\right)\cdot\left(1-\frac{L_M}{L}\right)\right]}\,. \qquad (37)$$

A rise in the degree of relative exposure reduces the wedge between insiders and outsiders, resulting in a decrease of the expected outside wage (see (36)). The wages of unskilled formal-sector workers fall, thereby, reducing both ω and θ. Since unskilled formal-sector workers become cheaper, $\frac{L_M}{L}$ expands (see Figure 6).

Figure 6: Simulation Results of the Elaborate Model 2



Source: Authors' calculations.

5.4 Introducing Structural Reforms into the Model

Having characterized the comparative statics of the three model versions, we can now turn to the distributional effects of structural reforms. Structural reform efforts of developing countries can be grouped into four categories: (a) product market deregulation, (b) labor market deregulation, (c) opening up to trade and foreign direct investment, and (d) privatization of public companies. Following Blanchard and Giavazzi (2001), these policy measures are integrated into the model in a highly abstract fashion. In the first step of the analysis, we discuss their impact on the model parameters μ, η, β, and ψ (see Matrix A of Figure 7). Linking the outcome of this exercise to the simulation results of Sections 5.2 and 5.3 (see Matrix B of Figure 7), we then derive the impact of structural reforms on wages and employment (see Matrix C of Figure 7).

Figure 7: Impact of Structural Reforms on Wages and Employment

Matrix A

	μ	η	β	ψ
DP		+		
DL			–	+
O	+	+		
P		?	–	+

x

Matrix B

	ω	θ	$\frac{L_M}{L}$
μ	–	?	+
η	?	–	+
β	+	+	–
ψ	–	–	+

=

Matrix C

	ω	θ	$\frac{L_M}{L}$
DP	?	–	+
DL	– –	– –	+ +
O	– ?	? –	+ +
P	? – –	? – –	? + +

DP = Deregulation of Product Markets, DL = Deregulation of Labor Markets, O = Opening Up to Trade and FDI, P = Privatization

By product market deregulation we mean policy reforms that increase the degree of competition in the modern sector, such as the reduction of administrative market-entry barriers and tax distortions, and the implementation of tougher antitrust enforcement. They are captured in the model by an increase in the absolute value of the elasticity of substitution in the formal sector, η.

Labor market deregulation has an impact on two model parameters. First, weakening extension agreements and closed-shop arrangements, restricting the right to strike, and other measures to curb unskilled workers' bargaining power are reflected in a reduction of β. Second, measures to reduce insider power, such as cutting the legal period of notice, and lowering redundancy payments and other administrative dismissal costs are modeled by an increase of the degree of relative exposure ψ.

Formally speaking, the model depicts a closed economy. Yet, it is still suitable to analyze the distributional effects of opening up to trade and foreign direct investment. Both the production of tradable goods and the inflow of foreign direct investment tend to be concentrated in the formal sector. As a result, a reduction of trade barriers should shift demand from the informal to the formal sector. In the model, this is reflected in a rise of μ. Additionally, opening up to foreign competition – be it via trade or via foreign direct investment – raises the degree of competition in the formal sector ($\eta\uparrow$).

Introducing privatization into the model is slightly more complex. This is because the impact of privatization on the degree of competition in product markets is ambiguous. On the one hand, Haskel and Szymanski (1993) argue that a shift from public to private ownership changes the objective function of the privatized entity. Public companies are thought to pursue the interests of all stakeholders, i.e., capital owners, workers, and consumers, while private firms confine themselves to profit maximization. Consequently, private firms are more likely to abuse market power than public companies. On the other hand, privatization is often accompanied by product market deregulation. This is done by replacing state monopolies by competitive market structures and by phasing out other types of administrative interference in the market. Furthermore, when balancing the interests of consumers and producers, regulators tend to favor producers in the case of public companies, but consumers in the case of private firms. Hence, antitrust rules tend to be more strictly enforced after privatization. For these reasons, the impact of privatization on η is undetermined. Privatization also often goes hand in hand with de-unionization and the weakening of job security. Both union density and co-determination are usually higher in public companies than in private firms. Consequently, privatization can be modeled as a reduction in β.[23] Furthermore, public employees frequently enjoy preferential treatment with respect to dismissal protection since

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23 See, for example, Haskel and Sanchis (1995).

(a) soft budget constraints in the public sector prevent mass layoffs in the first place, (b) the legal rules governing the dismissal of public employees are more stringent, and/or (c) their application is more strictly enforced. Hence, privatization should increase ψ.

In Sections 5.2 and 5.3, we already discussed the impact of changes of the parameters μ, η, β, and ψ on the relative wage of unskilled workers in the formal sector, on the sectoral wage premium of unskilled formal-sector workers, and on the unskilled employment share of the formal sector (see Matrix B of Figure 7). Referring to these results, we can now derive the wage and employment effects of structural reforms (see Matrix C of Figure 7).

Product market deregulation feeds into wages and employment via the rise in the degree of competition in the formal sector. Depending on the elasticity of substitution between skilled and unskilled workers in the formal sector, the relative wage of unskilled formal-sector workers may either increase or decrease. The effect on the other two endogenous variables, however, is unambiguous. The sectoral wage premium of unskilled formal-sector workers falls ($\theta\downarrow$), and the unskilled employment share of the formal sector rises ($\frac{L_M}{L}\uparrow$).

In the case of labor market deregulation, the distributional effects are clear cut. Due to the decline in union bargaining power and the rise in relative exposure, wages of unskilled formal-sector workers deteriorate, both relative to their skilled co-workers ($\omega\downarrow$) and relative to unskilled informal-sector workers ($\theta\downarrow$), and the unskilled employment share of the formal sector rises ($\frac{L_M}{L}\uparrow$).

There are two transmission mechanisms through which opening up to trade and foreign direct investment has an impact on wages and employment: both μ and η increase. The first effect reduces ω, has an indeterminate effect on θ, and raises $\frac{L_M}{L}$. The second effect has an indeterminate effect on ω, and causes θ to fall but $\frac{L_M}{L}$ to rise. Aggregating the two partial effects, it can be inferred that opening up to trade and foreign direct investment should increase the employment share of the formal sector ($\frac{L_M}{L}\uparrow$). Its impacts on the relative wage and the sectoral wage premium of unskilled formal-sector workers, however, are ambiguous.

Strictly speaking, analyzing the distributional effects of privatization does not render clear cut results either. This is because the impact of privatization on the degree of competition in the formal sector is undetermined. However, since the wage and employment effects of both the fall in union bargaining power and the rise in relative exposure point in the same direction, it is likely that privatization causes ω and θ to decrease and $\frac{L_M}{L}$ to increase.

5.5 Linking Empirical and Theoretical Results

To test the validity of our model for understanding the distributional effects of structural reforms in developing countries, we evaluate whether it helps explain the stylized facts on post-reform trends in Bolivian wage inequality. In Section 2, we saw that Bolivia has made considerable structural reform progress since 1985 (see Vector D of Figure 8). Product and labor markets were deregulated, the economy was opened up to trade and foreign direct investment, and public enterprises were privatized. Only labor market deregulation is still pending.

Linking the outcome of this analysis to the wage and employment effects of the individual policy reforms (see Matrix C of Figure 8), we can derive the distributional effects of the *Bolivian* structural reform process (see Vector E of Figure 8). According to our model, there should be two beneficiaries: skilled workers and unskilled ex-informal-sector workers. Both groups of workers can improve their income position relative to unskilled formal-sector workers ($\omega \downarrow$ and $\theta \downarrow$). Additionally, the rise of the unskilled employment share of the formal sector ($\frac{L_M}{L} \uparrow$) facilitates some unskilled ex-informal-sector workers to gain employment in the formal sector. This employment shift is rewarded by a pay rise of θ.

Figure 8: Distributional Effects of the Bolivian Structural Reform Process

Vector D

	DP	DL	O	P
Bolivia	+		+	+

x

Matrix C

	ω	θ	$\frac{L_M}{L}$
DP	?	–	+
DL	– –	– –	+ +
O	– ?	? –	+ +
P	? – –	? – –	? + +

=

Vector E

	ω	θ	$\frac{L_M}{L}$
Bolivia	? – ? ? – –	– ? – ? – –	+ + + ? + +

DP = Deregulation of Product Markets, DL = Deregulation of Labor Markets, O = Opening Up to Trade and FDI, P = Privatization

How do these model implications translate into the stylized facts outlined in Section 5.1? A rise in ω amplifies wage inequality, while a fall in θ and a rise in $\frac{L_M}{L}$ reduce wage inequality. As long as the first effect ($\omega\downarrow$) dominates the second effect ($\theta\downarrow$ and $\frac{L_M}{L}\uparrow$), wage inequality increases. Furthermore, due to the second effect, the rise in wage inequality is most pronounced at the upper tail of the distribution. In summary, our model can replicate stylized fact 1.

In our model, there are only two components contributing to wage inequality: (a) returns to skill, which are reflected in the wage differential between skilled and unskilled workers, and (b) rents earned by unskilled formal-sector workers. In the regression equation (8), returns to skill are captured by *Education.* By contrast, we argue that workers' ability to appropriate rents depends on personal characteristics, such as union membership, party affiliation or social status of the family, which can not be included in the empirical analysis for lack of disaggregate data. These omitted variables are, thus, only reflected in the *Residual.*

According to our model, the rise in wage inequality is due to declining rents for unskilled formal-sector workers. As a result, returns to skill become more important in wage setting, as evidenced by a rise in the correlation coefficient between *Education* and wages. Consequently, this variable is the main contributor to the rise in wage inequality. In the same vein, omitted variables become less important in wage setting, thereby counteracting wage inequality. This is in line with stylized facts 2 and 3.

6 Conclusion

The key message of the paper is that by introducing structural reforms into a rent-based dual-economy model, we can derive a theoretical explanation for post-reform trends in the wage inequality of developing countries. The scope for future research is wide. First, any empirical result derived by analyzing urban-only household surveys are subject to two well-known limitations. One is that nothing is known about the rural areas. The other is that even if we focus only on the urban areas, it is hard to control for the effects of rural-urban migration. Second, the present model takes a one-sided view of the informal sector. Following Fields' (1975) "staging hypothesis" we simply see it as a buffer for those workers who do not find one of the rationed formal-sector jobs. Empirical evidence, however, suggests that at least some informal-sector workers prefer their current employment status over formal-sector employment (Thomas 1992). In order to test the applicability of the "staging hypothesis" for the Bolivian labor market, it would be necessary to come up with a theoretically consistent and empirically implementable concept of the formal-informal sector dichotomy.

Appendix A: Regression Results

Table A1 — Encuesta Continua de Hogares 1989

Variables	β_j	μ_{Z_j}	σ_{Z_j}	cor(Z_j, ln(yph))	S_j	S_j
c	-0.5303	—	—	—	—	—
age	0.0695	35.81	11.68	0.1320	0.1207	—
$age^2/10$	-0.0073	141.86	91.17	0.0993	-0.0748	0.0460
school	0.0624	8.94	4.66	0.3487	0.1142	0.1142
gender	-0.1433	0.39	0.49	-0.1487	0.0117	0.0117
prod. worker	-0.7294	0.12	0.32	-0.1229	0.0324	—
non-prod. worker	-0.5916	0.42	0.49	0.1048	-0.0344	—
employer	0.0292	0.03	0.18	0.1624	0.0010	—
self-employed	-0.4237	0.42	0.49	-0.1160	0.0273	—
ind. profes. (dropped)	—	—	—	—	—	0.0263
agriculture	-0.2872	0.02	0.13	0.0366	-0.0015	—
mining	-0.2560	0.02	0.14	-0.0095	0.0004	—
manufacturing	-0.3362	0.13	0.34	-0.0423	0.0054	—
utilities	0.0284	0.01	0.07	0.0386	0.0001	—
construction	-0.2405	0.07	0.26	0.0045	-0.0003	—
trade & commerce	-0.5215	0.25	0.43	-0.1947	0.0496	—
hotels & restaurants	-0.3572	0.04	0.21	-0.0481	0.0040	—
transport	-0.2021	0.09	0.28	0.0627	-0.0040	—
education (not available)	—	—	—	—	—	—
other services	-0.2646	0.26	0.44	0.1051	-0.0138	—
public admin.	-0.2127	0.08	0.26	0.0594	-0.0038	—
fin.&busi.serv (dropped).	—	—	—	—	—	0.0361
chuqisaca	0.3637	0.04	0.19	0.0059	0.0005	—
la paz	0.2907	0.43	0.50	-0.1212	-0.0197	—
cochabamba	0.4148	0.17	0.37	0.0657	0.0115	—
oruro	0.1403	0.07	0.25	-0.0633	-0.0025	—
tarija	0.4022	0.03	0.16	-0.0021	-0.0002	—
santa cruz	0.6979	0.21	0.41	0.1543	0.0493	—
beni	0.5550	0.02	0.13	0.0266	0.0022	—
pando	0.7768	0.00	0.05	0.0243	0.0012	—
potosi (dropped)	—	—	—	—	—	0.0423
sum = R^2	—	—	—	—	0.2765	0.2765
nobs = 5186 R^2 = 0.2764	$\mu_{\ln(yph)}$ = 0.9776		$\sigma_{\ln(yph)}$ = 0.8879			

Table A2 — Encuesta Continua de Hogares 1991

Variables	β_j	μ_{Z_j}	σ_{Z_j}	cor(Z_j, ln(yph))	S_j	S_j
c	-0.7697	—	—	—	—	—
age	0.0641	34.76	11.45	0.1668	0.1333	—
$age^2/10$	-0.0063	133.91	87.20	0.1350	-0.0804	0.0530
school	0.0706	9.30	4.45	0.3711	0.1269	0.1269
gender	-0.2127	0.37	0.48	-0.1541	0.0172	0.0172
prod. worker	-0.5564	0.20	0.40	-0.1501	0.0361	—
non-prod. worker	-0.4625	0.38	0.49	0.1683	-0.0412	—
employer	0.0422	0.05	0.21	0.1568	0.0015	—
self-employed	-0.3473	0.36	0.48	-0.1376	0.0250	—
ind. profes. (dropped)	—	—	—	—	—	0.0214
agriculture	-0.4895	0.01	0.12	-0.0231	0.0015	—
mining	0.0203	0.02	0.15	0.0409	0.0001	—
manufacturing	-0.2977	0.19	0.39	-0.0697	0.0089	—
utilities	0.0181	0.01	0.09	0.0315	0.0001	—
construction	-0.1978	0.10	0.30	-0.0176	0.0011	—
trade & commerce	-0.3778	0.25	0.43	-0.1598	0.0284	—
hotels & restaurants	-0.3232	0.04	0.20	-0.0666	0.0048	—
transport	-0.1739	0.08	0.27	0.0472	-0.0024	—
education	-0.0871	0.08	0.28	0.1366	-0.0036	—
other services	-0.0980	0.10	0.30	0.0577	-0.0019	—
public admin.	-0.1185	0.06	0.24	0.0724	-0.0023	—
fin.&busi.serv (dropped).	—	—	—	—	—	0.0347
chuqisaca	0.3425	0.04	0.20	-0.0040	-0.0003	—
la paz	0.2687	0.42	0.49	-0.0853	-0.0123	—
cochabamba	0.3569	0.14	0.35	0.0374	0.0050	—
oruro	0.1641	0.06	0.23	-0.0309	-0.0013	—
tarija	0.4005	0.03	0.17	0.0103	0.0008	—
santa cruz	0.5564	0.26	0.44	0.1015	0.0269	—
beni	0.5748	0.02	0.14	0.0332	0.0029	—
pando (not available)	—	0	0	—	0	—
potosi (dropped)	—	—	—	—	—	0.0217
sum = R^2	—	—	—	—	0.2749	0.2749
nobs = 8157 R^2 = 0.2745 $\mu_{ln(yph)}$ = 0.9080 $\sigma_{ln(yph)}$ = 0.9189						

Table A3 —Encuesta Continua de Hogares 1993

Variables	β_j	μ_{Z_j}	σ_{Z_j}	cor(Z_j, ln(yph))	S_j	S_j
c	-1.0998	—	—	—	—	—
age	0.0699	34.94	11.57	0.1333	0.1134	—
$age^2/10$	-0.0072	135.48	88.13	0.0953	-0.0633	0.0501
school	0.0792	9.84	5.31	0.5108	0.2260	0.2260
gender	-0.1892	0.38	0.48	-0.1739	0.0168	0.0168
prod. worker	-0.3486	0.19	0.40	-0.1460	0.0212	—
non-prod. worker	-0.2728	0.39	0.49	0.2242	-0.0314	—
employer	0.3368	0.07	0.25	0.2341	0.0212	—
self-employed	-0.3367	0.33	0.47	-0.2683	0.0449	—
ind. profes. (dropped)	—	—	—	—	—	0.0558
agriculture	-0.2068	0.02	0.13	-0.0103	0.0003	—
mining	0.1352	0.02	0.12	0.0448	0.0008	—
manufacturing	-0.2689	0.19	0.39	-0.0963	0.0107	—
utilities	0.2154	0.00	0.07	0.0475	0.0007	—
construction	-0.1102	0.10	0.29	-0.0158	0.0005	—
trade & commerce	-0.3314	0.25	0.44	-0.1969	0.0299	—
hotels & restaurants	-0.3606	0.05	0.22	-0.1044	0.0087	—
transport	-0.1174	0.10	0.30	0.0391	-0.0014	—
education	-0.2587	0.03	0.17	0.0790	-0.0037	—
other services	-0.0503	0.14	0.34	0.1564	-0.0028	—
public admin.	0.1275	0.06	0.24	0.1289	0.0041	—
fin.&busi.serv (dropped).	—	—	—	—	—	0.0477
chuqisaca	0.1654	0.04	0.19	-0.0297	-0.0010	—
la paz	0.2636	0.42	0.49	-0.0529	-0.0072	—
cochabamba	0.3155	0.14	0.35	0.0161	0.0018	—
oruro	-0.1066	0.06	0.23	-0.0865	0.0023	—
tarija	0.2443	0.03	0.18	-0.0294	-0.0013	—
santa cruz	0.5972	0.26	0.44	0.1390	0.0384	—
beni	0.4966	0.02	0.14	0.0142	0.0010	—
pando (not available)	—	0	0	—	0	—
potosi (dropped)	—	—	—	—	—	0.0339
sum = R^2	—	—	—	—	0.4303	0.4303
nobs = 5917 $R^2 = 0.4303$ $\mu_{\ln(yph)} = 0.9624$ $\sigma_{\ln(yph)} = 0.9503$						

Table A4 — Encuesta Continua de Hogares 1995

Variables	β_j	μ_{Z_j}	σ_{Z_j}	cor(Z_j, ln(yph))	S_j	S_j
c	-0.4607	—	—	—	—	—
age	0.0489	34.73	11.44	0.1663	0.0982	
age^2/10	-0.0047	133.69	86.58	0.1343	-0.0576	0.0406
school	0.0646	10.24	5.68	0.4199	0.1622	0.1622
gender	-0.1687	0.38	0.49	-0.1189	0.0103	0.0103
prod. worker	-0.4793	0.19	0.39	-0.1725	0.0343	—
non-prod. worker	-0.4258	0.36	0.48	0.1413	-0.0304	—
employer	0.1571	0.09	0.29	0.2130	0.0102	—
self-employed	-0.2672	0.35	0.48	-0.1593	0.0214	—
ind. profes. (dropped)	—	—	—	—	—	0.0354
agriculture	-0.5061	0.02	0.14	-0.0306	0.0023	—
mining	0.1289	0.02	0.13	0.0474	0.0008	—
manufacturing	-0.3655	0.19	0.39	-0.1144	0.0173	—
utilities	0.1482	0.00	0.07	0.0410	0.0004	—
construction	-0.2601	0.10	0.30	-0.0271	0.0022	—
trade & commerce	-0.3398	0.27	0.44	-0.1116	0.0177	—
hotels & restaurants	-0.2542	0.05	0.22	-0.0554	0.0033	—
transport	-0.2366	0.09	0.28	0.0097	-0.0007	—
education	-0.3303	0.03	0.16	0.0448	-0.0025	—
other services	-0.0426	0.14	0.35	0.1527	-0.0024	—
public admin.	-0.0171	0.05	0.22	0.0771	-0.0003	—
fin.&busi.serv (dropped).	—	—	—	—	—	0.0383
chuqisaca	0.2561	0.04	0.19	-0.0406	-0.0021	—
la paz	0.3221	0.41	0.49	-0.0713	-0.0119	—
cochabamba	0.3656	0.13	0.34	0.0186	0.0024	—
oruro	0.1558	0.06	0.24	-0.0612	-0.0024	—
tarija	0.3130	0.03	0.18	-0.0260	-0.0015	—
santa cruz	0.7156	0.27	0.45	0.1460	0.0491	—
beni	0.6150	0.02	0.13	0.0267	0.0023	—
pando (not available)	—	0	0	—	0	—
potosi (dropped)	—	—	—	—	—	0.0359
sum = R^2	—	—	—	—	0.3227	0.3227
nobs = 8037 $R^2 = 0.3231$	$\mu_{ln(yph)} = 1.0608$		$\sigma_{ln(yph)} = 0.9485$			

Table A5 — Encuesta Nacional de Empleo 1997

Variables	β_j	μ_{Z_j}	σ_{Z_j}	cor(Z_j, ln(yph))	S_j	S_j
c	-0.5594	—	—	—	—	—
age	0.0519	36.22	11.82	0.1582	0.1009	—
$age^2/10$	-0.0049	145.18	90.90	0.1263	-0.0590	0.0419
school	0.0673	10.72	5.84	0.4800	0.1961	0.1961
gender	-0.1803	0.39	0.49	-0.1149	0.0105	0.0105
prod. worker	-0.2514	0.16	0.37	-0.1845	0.0177	—
non-prod. worker	-0.1222	0.37	0.48	0.2224	-0.0137	—
employer	0.4604	0.08	0.26	0.2066	0.0261	—
self-employed	-0.0953	0.38	0.49	-0.2192	0.0105	—
ind. profes. (dropped)	—	—	—	—	—	0.0407
agriculture	-0.5892	0.02	0.15	-0.0913	0.0084	—
mining	0.1224	0.01	0.12	0.0365	0.0005	—
manufacturing	-0.3062	0.20	0.40	-0.1205	0.0153	—
utilities	0.3136	0.01	0.09	0.0740	0.0021	—
construction	-0.1666	0.10	0.30	-0.0227	0.0012	—
trade & commerce	-0.3277	0.23	0.42	-0.1496	0.0216	—
hotels & restaurants	-0.1494	0.05	0.22	-0.0557	0.0019	—
transport	-0.2215	0.10	0.30	0.0145	-0.0010	—
education	-0.1195	0.06	0.24	0.1457	-0.0043	—
other services	-0.0301	0.12	0.32	0.1251	-0.0013	—
public admin.	-0.0254	0.05	0.21	0.0985	-0.0005	—
fin.&busi.serv (dropped).	—	—	—	—	—	0.0439
chuqisaca	0.0828	0.04	0.21	-0.0399	-0.0007	—
la paz	0.1812	0.37	0.48	-0.1064	-0.0097	—
cochabamba	0.3297	0.16	0.37	0.0668	0.0085	—
oruro	0.1034	0.06	0.23	-0.0204	-0.0005	—
tarija	0.0764	0.03	0.18	-0.0755	-0.0011	—
santa cruz	0.4931	0.28	0.45	0.1181	0.0271	—
beni	0.3364	0.02	0.14	0.0089	0.0004	—
pando	0.7588	0.01	0.07	0.0376	0.0022	—
potosi (dropped)	—	—	—	—	—	0.0262
sum = R^2	—	—	—	—	0.3593	0.3593
nobs = 4549 R^2 = 0.3593 $\mu_{\ln(yph)}$ = 1.2421 $\sigma_{\ln(yph)}$ = 0.9618						

Appendix B: The Model

Appendix B1

The negotiation partners solve the following program

$$
\begin{aligned}
\max_{\{W_{M_j}\}} \quad & \Omega_{V_j} = \Gamma_{U_j}^{\beta_j}\, \Gamma_{M_j}^{1-\beta_j}, \\
\text{s.t.} \quad & \Gamma_{U_j} = L_{M_j} \cdot \left(W_{M_j} - Z\right), \\
& \Gamma_{M_j} = P_{Mj} \cdot {L_{M_j}}^{\alpha} \cdot {H_{M_j}^{v}}^{1-\alpha} - W_{M_j} \cdot L_{M_j} - Q_{M_j} \cdot H_{M_j}^{v}.
\end{aligned}
\tag{B1}
$$

The first order condition for the optimal wage reads

$$
\varepsilon_{\Gamma_{U_j}, W_{M_j}} = -\frac{1-\beta_j}{\beta_j} \cdot \varepsilon_{\Gamma_{M_j}, W_{M_j}} \tag{B2}
$$

with

$$
\varepsilon_{\Gamma_{U_j}, W_{M_j}} = \frac{W_{M_j}}{W_{M_j} - Z} - 1 - \alpha \cdot (\eta - 1), \tag{B3}
$$

$$
\varepsilon_{\Gamma_{M_j}, W_{M_j}} = -\alpha \cdot (\eta - 1) \tag{B4}
$$

being the wage elasticities of the unions' and the firms' stake in the negotiations. Substituting (B3) and (B4) into (B2), we obtain

$$
W_{M_j} = \left[1 + \frac{\beta}{\alpha \cdot (\eta - 1)}\right] \cdot Z. \tag{24}
$$

Appendix B2

It is straightforward to show that in a symmetric equilibrium, wages of skilled and unskilled workers, Q_{M_j} and W_{M_j}, prices and quantities demanded of the modern good, P_{M_j} and M_j, and the quantities demanded of skilled and unskilled workers, $H^v_{M_j}$, $H^f_{M_j}$ and L_{M_j}, are equal for all formal-sector firms.

Inserting (16) and (17) into the zero-profit condition of the formal sector, we obtain the share of skilled workers who are employed as production workers

$$\frac{H^v_M}{H} = \frac{(1-\alpha)\cdot(\eta-1)}{\alpha+\eta\cdot(1-\alpha)}. \tag{B5}$$

Dividing (16) by (17) and using (B5), we arrive at equation (27)

$$\omega := \frac{W_M}{Q_M} = \frac{\alpha\cdot(\eta-1)}{\eta-\alpha\cdot(\eta-1)}\cdot\frac{H}{L}\cdot\left(\frac{L_M}{L}\right)^{-1}. \tag{27}$$

Inserting (10) and (B5) into (16), we can solve for equation (28)

$$1+\theta := \frac{W_M}{W_T} = \frac{\mu}{1-\mu}\cdot\frac{\eta-1}{\eta}\cdot\alpha\cdot\frac{1-\frac{L_M}{L}}{\frac{L_M}{L}}. \tag{28}$$

Combining (24) and (26) yields equation (29)

$$1+\theta := \frac{W_M}{W_T} = \frac{\left[1+\frac{\beta}{\alpha\cdot(\eta-1)}\right]\cdot\left(1-\frac{L_M}{L}\right)}{1-\left[1+\frac{\beta}{\alpha\cdot(\eta-1)}\right]\cdot\frac{L_M}{L}}. \tag{29}$$

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