# Inflation and Labor Market Flexibility: The Squeaky Wheel Gets the Grease

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## Abstract

Inflation can "grease" the wheels of the labor market by relaxing downward wage rigidity but it can also increase uncertainty and have a negative "sand" effect. This paper studies the grease effect of inflation by looking at whether the interaction between inflation and labor market regulations affects how employment responds to changes in output. The results show that in industrial countries with highly regulated labor markets, the grease effect of inflation dominates the sand effect. In the case of developing countries, we rarely find a significant effect of inflation or labor market regulations and provide evidence indicating that this could be due to the presence of a large informal sector and limited enforcement of *de jure* labor market regulations.

**JEL Codes**: E24; E31; E52
**Keywords**: Employment; Unemployment; Flexibility; Inflation; Deflation; Job Security

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# 1. Introduction

In his 1972 Presidential Address to the American Economic Association, James Tobin stated: "Unemployment and inflation still preoccupy and perplex economists, statesmen, journalists, housewives, and everyone else. …The connection between them is… the major area of controversy and ignorance in macroeconomics." In the same paper, Tobin suggested "No one has devised a way of controlling average wage rates without intervening in the competitive struggle over relative wages. Inflation lets this struggle proceed and blindly, impartially, impersonally, and nonpolitically scales down all its outcomes" (Tobin, 1972 p.13). In other words, inflation may play a beneficial role by adding grease to the wheels of the labor market.[1] Five years later, Milton Friedman's Nobel Lecture focused on the sand effects of inflation. According to the sand view, inflation increases uncertainty and, by arbitrarily changing relative prices and wages, leads to resource misallocation and lower levels of employment (Friedman, 1977).

The empirical evidence has not been kind to the grease hypothesis. Akerlof, Dickens, and Perry (1996, 2000) find some evidence in support of the idea that when inflation is below 1.5 percent there is a long-run tradeoff between inflation and unemployment. Card and Hyslop (1996) find evidence that an increase in inflation allows real wages to fall faster, but they find no evidence that inflation affects wage adjustment across local labor markets. Groshen and Schweitzer (1997) use firm-level data to distinguish the grease effect of inflation from its sand effect. They find that while inflation below 5 percent has a positive but not statistically significant grease effect, inflation above 5 percent has a statistically significant sand effect.

The main claim of this paper is that the lack of success in identifying the grease effect of inflation is due to the focus on the US labor market,[2] which, being among the most flexible in the world, does not need much grease to start with. In fact, one would expect that grease effects should be more important in the highly regulated European labor market than in the fairly flexible US market. We tackle this issue by looking at whether the interaction between inflation and labor market regulations affects how employment responds to changes in output (the employment Okun coefficient). We find strong evidence that in industrial countries with highly regulated labor markets, inflation reduces the sensitivity of employment to changes in output. We also find some evidence in support of the idea that lower employment elasticity is driven by the fact that inflation increases real wage flexibility. We conclude that

[1] The grease hypothesis suggests that inflation can speed the adjustment to the long run equilibrium but is consistent with the idea of a vertical long-run Phillips Curve. A second class of models rejects the idea of a vertical long-run Phillips Curve and, by using near-rational wage setting behavior, shows that at low levels of inflation there is a long-run trade-off between inflation and unemployment (Akerlof, Dickens, and Perry, 2000).

[2] A notable exception is Decressin and Decressin (2002). They use individual-level data for Germany to evaluate the grease and sand of effect of inflation. Their results are similar to those of Groshen and Schweitzer (1997), and they conclude that inflation does not weaken the macroeconomic effects of labor market regulations.

in industrial countries with highly regulated labor markets, the grease effect of inflation dominates the sand effect. We find that the opposite is true for industrial countries that are characterized by more flexible labor markets. In this set of countries (which includes the United States), we find that inflation increases employment elasticity and, thus, the sand effect of inflation dominates the grease effect. This suggests that inflation does grease the wheels of the labor market, but only those that squeak the most.

Looking at developing countries, we rarely find a statistically significant effect of inflation or labor market regulations. We posit that this may be because most developing countries do not enforce regulations. We present some evidence in support of this hypothesis by showing that the effect of inflation and labor market regulations is higher in countries characterized by higher levels of rule of law.

Three papers that are closely related to our work are Ball (1997), Wyplosz (2001), and González (2002). The first paper studies the disinflation process in 20 OECD countries and shows that disinflation was associated with an increase in the natural rate of unemployment. Ball (1997) also shows that the effect of the disinflation process was larger in countries characterized by a highly regulated labor market. Wyplosz (2001) recognizes that the grease and sand effect of inflation may vary with the degree of labor market rigidities and studies the cases of Germany, France, the Netherlands, and Switzerland. His results differ from those of Groshen and Schweitzer (1997) because he finds a sand effect at very low levels of inflation and a grease effect at higher levels of inflation. González (2002) computes employment and unemployment Okun coefficients for a large set of Latin American countries over the 1970-1996 period and discusses how structural reforms and the disinflation process may have affected how employment and unemployment respond to output shocks. Contrary to our work, however, González does not test formally the presence of a relationship among employment elasticity, labor market regulation, and inflation. Finally, this paper is also related to Blanchard (1999) and Blanchard and Wolfer's (2000) work emphasizing the importance of the interaction between economic shocks and labor market institutions.

The rest of the paper is organized as follows. Section 2 discusses a highly stylized model that relates employment to inflation and labor market regulations. Section 3 presents the empirical evidence on the determinants of employment elasticity. Section 4 looks at wage rigidity. Section 5 concludes.

## 2. A Simple Model

We set the stage for the empirical analysis by discussing an extremely stylized model that focuses on the interaction between inflation and labor market regulations. This model is a basic extension of Bertola

(1990) and has no pretense of originality.[3] However, we think it useful to clarify the ideas and provide a clear set of testable hypotheses.

Bertola (1990) studies the problem of a risk-neutral representative firm that chooses employment in order to maximize the present value of expected profits:

$$\underset{\{L_i\}}{Max}\, E_t \left\{ \sum_{i=0}^{\infty} \left( \frac{1}{1+r} \right)^i \left[ R(Z_{t+i}, L_{t+i}) - W_{t+i} L_{t+i} - C(L_{t+i} - L_{t+i-1}) \right] \right\} \tag{1}$$

He defines the process $\{Z_t\}$ as an index of business conditions and $R(Z_{t+i}, L_{t+i})$ as the operating revenues obtained by employing *L* homogenous workers. The function $R(Z_{t+i}, L_{t+i})$ is assumed to be increasing and concave in L, with $R(Z_{t+i}, 0)$ =0. All the variables are described in real terms and the wage process $\{W_t\}$ is assumed to be exogenous and will be described later. The firm faces firing costs described by the following function:[4]

$$C(L_{t+i} - L_{t+i-1}) = \begin{cases} 0 & if \quad L_{t+i} - L_{t+i-1} > 0 \\ -F(\rho)(L_{t+i} - L_{t+i-1}) & if \quad L_{t+i} - L_{t+i-1} < 0 \end{cases} \tag{2}$$

We assume that firing costs *F* depend on an index of labor market regulations $\rho \in (0,1)$, with *F'>0* and *F(0)=0*. Changes in business conditions are modeled using a two-state Markov chain. The economy moves from a "good" state of the world $(Z_g)$ to a "bad" $(Z_b < Z_g)$ state of the world with probability $1 - P_g$, and moves from a bad state of the world to a good state of the world with probability $1 - P_b$. To simplify the analysis, we assume that $P_b$ =0. This is equivalent to assuming that bad states of the world only last one period. Following Bertola, we assume that the firm initiates all separations and that desired employment is higher in good times. Formally: $R_Z > 0$, and $R_{ZL} > 0$.

In the absence of nominal rigidities, real wages would be exogenously set to $W_g$ in good states of the world, and set equal to the reservation wage ($\overline{W}$) in bad states of the world ($\overline{W} < W_g$). It is assumed that the differential between $W_g$ and the reservation wage is smaller than the one that would lead to a

[3] Note for the referees: The model could be dropped or moved to the Appendix.
[4] Bertola (1990) assumes both firing and hiring costs.

lower level of employment in the good state of the world (therefore, it is always true that $L_b \leq L_g$). We now depart from Bertola (1990) and assume nominal wage rigidities (this is the only significant departure with respect to Bertola's paper). We model nominal rigidities in a rather brutal way and assume that in bad states of the world real wages are given by:

$$W_{b,t} = \rho W_{t-1}(1-\pi_t) + (1-\rho)\overline{W} \tag{3}$$

where $\pi$ is inflation and $W_{t-1}$ is the real wage in the previous period.[5] In highly regulated labor markets ($\rho$=1), nominal wages in period *t* are equal to nominal wages in period *t-1*. Therefore, real wages in period *t* are equal to real wages in the previous period minus inflation in period *t*. In labor markets with no rigidities, bad time real wages are set equal to the reservation wage $\overline{W}$.[6] Clearly the above equation relies on the strong assumption that either workers only care about nominal wage and not real wages (if not indexation would arise) or that they are myopic and always expect inflation to be equal to zero. So, Equation 3 does not allow any room for indexation (either to past or current inflation).

As we assumed that bad states of the world last only one period, we can rewrite Equation 3 as:

$$W_{b,t} = \rho W_g(1-\pi_t) + (1-\rho)\overline{W} \tag{3'}$$

At this point, it is important to note that Equations 2 and 3 assume that the index of labor market regulations affects both firing costs and wage flexibility. Bertola and Rogerson (1997) provide a rationale for such an assumption. They point out that without wage rigidities, job protection makes little sense because entrepreneurs would have the option to drive real wages close to zero and thus make job protection irrelevant. The same would apply to a situation in which entrepreneurs cannot touch real wages but can fire at will. It is therefore natural that the political and economic institutions that lead to a high level of job protection will also lead to higher wage rigidity.[7] We use data from Botero et al. (2003) to

[5] To be more precise, we should set $W_t^b = Max\left[\rho W_{t-1}(1-\pi_t) + (1-\rho)\overline{W}, \overline{W}\right]$. To simplify things, we will assume that $\left[\rho W_{t-1}(1-\pi_t) + (1-\rho)\overline{W}\right] \geq \overline{W}$. This is equivalent to assuming that $W_{t-1}(1-\pi_t) \geq \overline{W}$.

[6] We assume no correlation between $\overline{W}$ and $\rho$. However, the reservation wage is likely to be affected by factors like unemployment insurance, which in turn could be correlated with the presence of labor market regulations.

[7] Bertola and Rogerson (1997) state: "the apparent association of wage equalization and job security provisions can be intuitively rationalized in terms of simple politico-economic considerations. When implemented in isolation, neither wage compression nor dismissal restrictions can fulfill a likely aim of intervention in the labor market—namely stabilization of labor incomes in the face of idiosyncratic (yet uninsurable) labor-demand shocks (p. 1169).

check whether there is empirical support for a positive correlation between the institutional determinants of wage rigidity and firing costs. In particular, we look at the correlation between their index of job security and their index of industrial relation laws. The latter measures, among other things, the presence and extent of a collective bargain system and the regulation of collective disputes. These factors should in turn proxy for the power of unions and hence for the institutional determinants of wage rigidity. Figure 1 indicates that there is a strong correlation between the two variables.

We solve the model using the same procedure used by Bertola (1990). Define the marginal revenue product of labor as: $M(Z,L) \equiv R_L$, and the dynamic shadow value of labor at time *t* as:

$$S_t \equiv E_t \left\{ \sum_{i=0}^{\infty} \left( \frac{1}{1+r} \right)^i \left[ M(Z_{t+i}, L_{t+i}) - W_{t+i} \right] \right\} \tag{4}$$

Assuming that the state of the world is observed before setting $L_t$, we have the following set of first-order conditions:

$$-F(\rho) \leq S_t \leq 0 \tag{5}$$

$$S_t = 0 \quad if \quad L_t > L_{t-1} \tag{6}$$

$$S_t = -F(\rho) \quad if \quad L_t < L_{t-1} \tag{7}$$

Labor demand is defined by a pair of employment levels $L_g \geq L_b$ that satisfy the first-order conditions in 5, 6, and 7. As quits are ruled out, labor demand also defines total employment. The latter will only decrease when the condition switches from good to bad, and increase when the condition switches from bad to good. In the presence of high firing costs or high wage flexibility, the firm may decide not to hire or fire. In this case employment will be constant across states of the world. For the sake of simplicity, we rule out this possibility and restrict our analysis to the case in which $L_g > L_b$. By substituting the definition of $S_t$ in the first order condition for good times, using the law of iterated expectations, and noting that $E_t(S_{t+1}) = P_g \times 0 + (1 - P_g) \times (\text{-F})$, it is easy to derive the following equation:

$$M(Z_g, L_g) = W_g + \frac{F(\rho)}{1+r}(1 - P_g) \tag{8}$$

Equation 8 implicitly defines labor demand in good times. It shows that positive firing costs cause good time wages to be lower than the value of the marginal product of labor. The concavity of *R* implies that employment is decreasing in *M(Z,L)*. Therefore, the presence of firing costs leads to less hiring during good times. Recalling that we assumed $P_b$=0 (hence, $E_t(S_{t+1})$=0), we can use the same procedure and derive the equation that implicitly defines labor demand in bad times:

$$M(Z_b, L_b) = \rho\left(W_g(1-\pi) - \overline{W}\right) + \overline{W} - F(\rho) \tag{9}$$

With positive firing costs, bad time wages are above the marginal product of labor (with no rigidities, the marginal product of labor would be equal to the reservation wage). The effect of labor market regulations on employment during bad times is not clear. On the one hand, they increase firing costs and lead to higher employment during bad times. On the other hand, they reduce wage flexibility and keep wages above the reservation wages and thus lead to less employment.

In contrast, the effect of inflation is clear. It always increases wage flexibility and therefore leads to more employment (with respect to a situation with lower inflation) during bad times. In this sense, the model does not predict any sand effect of inflation and only allows for a grease effect. A sand effect of inflation could be introduced by making marginal revenues negatively depend on inflation.

Figure 2 summarizes the main finding of the model. By increasing firing costs, labor market regulations lower labor demand (leading to lower employment) during good times. For the same reason, they increase labor demand during bad times. This positive effect on employment is counterbalanced by the fact that labor market regulations reduce wage flexibility and, by keeping wages above the reservation wage, may reduce employment. The overall effect on employment in bad times is therefore uncertain.

A clear implication of the model is that the effect of inflation is amplified by the presence of labor market regulations. In fact, if we were to assume that the labor market is perfectly flexible (i.e., $\rho$ *=0*), inflation would completely drop from the equations that determine labor demand.[8]

## 3. Estimation

The key message of the model of Section 2 is that inflation plays a useful grease role only in highly regulated labor markets. In this section, we estimate how the interaction between inflation and labor

[8] This can also be seen by computing the derivative: $\dfrac{\partial(-M_t(Z_b, L_b))}{\partial\pi} = \rho W_g$ .

market regulations affects the sensitivity of employment with respect to output. In particular, we will focus on the employment Okun coefficient (defined as the change in employment brought about by a change in output). We expect the Okun coefficient to be low when most of the adjustment to an output shock goes through a change in wages, and we expect the Okun coefficient to be high when most of the adjustment goes through employment. Within the framework of the model of Section 2, we can write the Okun coefficient as:

$$L_t - L_{t-1} = L(M(Z_t)) - L(M(Z_{t-1})) \tag{10}$$

where $L$ is a labor demand function and, by concavity of $R$, $L'<0$. By substituting Equations 8 and 9 into Equation 10, we can rewrite the Okun coefficient as a function of the exogenous variables:

$$\frac{\Delta L}{\Delta Z} = G\left(W_g, \overline{W}, \pi, \rho, r, P_g\right) \tag{11}$$

The main prediction of the model is that $G_{\pi,\rho} < 0$. We test this hypothesis by using the following specification:

$$\begin{aligned} DE_{i,t} &= (a_1 + a_2 INF_{i,t} + a_3 REG_{i,t} + a_4 REG_{i,t} \times INF_{i,t})DY + b_1 INF_{i,t} + b_2 REG_{i,t} + \\ &+ b_3 REG_{i,t} \times INF_{i,t} + cDE_{t-1,i} + \alpha_i + tYEAR + \varepsilon_{i,t} \end{aligned} \tag{12}$$

where $DE$ measures employment growth or its deviation (in percentage terms) from a log-linear trend (the deviation with respect to a HP trend yields similar results). We focus our analysis on employment (and not unemployment) because employment is measured more accurately and it is less affected by labor market participation decisions. In the robustness analysis, we will show that our results are robust to substituting employment with unemployment. $DY$ is GDP growth (or its deviation from a log linear trend).[9] INF is inflation, REG an index of labor market regulations, $\alpha$ a country fixed effect, and YEAR a time trend.[10] The parameters in parenthesis ($a_2$, $a_3$, and $a_4$) tell us how inflation, labor market regulations, and the interaction between the two affect employment elasticity. While the model of the previous section predicts $a_2$ to be negative, we already pointed out that the model could be modified (by introducing a sand effect) to make the sign of $a_2$ uncertain. Hence, we do not have a clear prediction for

[9] We use growth rate or deviation from trend because both employment and GDP are highly persistent.

the sign of $a_2$. We also do not have a clear prediction for $a_3$. In fact, labor market regulations could either increase (through their effect on wages) or decrease (thorough their effect on firing costs) employment elasticity. Our main parameter of interest is the one that measures the effect of the interaction between inflation and labor market regulations ($a_4$). In this case, the model yields a clear prediction and we expect $a_4$ to be negative, indicating that the grease effect of inflation is higher in countries with highly regulated labor markets. We do not have a clear prediction on the other variables that are introduced mainly as controls.

In order to estimate Equation 12, we need to identify a good proxy for REG. We measure labor market regulations by using an updated version of the job security index compiled by Pagés (2002) and based on Heckman and Pagés (2000) and Pagés and Montenegro (1999). The index of job security captures the marginal cost of dismissing a full-time worker with an open-ended contract. While this is not a perfect measure of job security, to the best of our knowledge it is the only available panel data set of the stringency on labor market regulations.[11] The original index measures dismissal costs in terms of monthly wages and ranges from zero (for the US) to 6.9 (for Venezuela until 1996). The industrial countries with the highest values of the index are Spain and Italy (with values that range between 3.2 and 3.8). We normalize the index so that it ranges between zero and one. The average value for all countries is 0.3, the average for industrial countries is 0.2 and the average for developing countries is 0.5 (see Table 1). Because the job security index derived by Heckman and Pagés focuses on dismissal costs, it should be clear that in order to use it to test our model we need to follow Bertola and Rogerson (1997) and assume that there is a set of common factors that determines both dismissal costs and nominal wage rigidity.

We estimate Equation 12 using two different panels that contain annual observations over the 1982-2000 period (in the robustness analysis, we also reproduce the results using a panel where all variables are averaged over seven three-year periods). The first panel focuses on industrial countries and the second on developing countries. In the sample of developing countries, we drop all observations for which inflation is above 30 percent (the results do not change if we use other thresholds or include all observations). Table 1 reports summary statistics for the variables used in the regression; the data sources are described in the Appendix.[12]

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[10] Using year fixed effects yield similar results.

[11] Botero et al. (2003) cover a larger sample of countries but their data set is only cross-sectional. At the same time, the data set compiled by Nickell et al. (2001) is of a panel nature but does not cover developing countries. There are two problems with the Heckman and Pagés (2000) index. On the one hand, the index may overestimate the true marginal cost of dismissals because it does not measure the dismissal cost of temporary workers. On the other hand, the index may underestimate the true marginal cost of dismissals because it does not measure the legal costs that could arise if the worker challenges the dismissal.

[12] The panel is unbalanced. In the sample of industrial countries, we have observations for the whole period (1982-2000) for Australia, Austria, Belgium, Canada, Finland, France, Italy, Japan, Norway, Switzerland, United Kingdom

We start the analysis by running a set of standard fixed effects regressions. Next, we look at possible problems with the estimation technique by running random effects estimations, checking whether the results are driven by outliers, and correcting for the bias introduced by the presence of the lagged dependent variable. Then, we address the problem of reverse causality by running instrumental variable regressions. Finally, we check whether our results are robust to the use of alternative measures of job security.

### *Evidence from Industrialized Countries*

Table 2 reports the results for industrial countries. Column 1 reports results for a standard fixed effects regression. It shows that the coefficient attached to *DYINF* is positive and statistically significant, indicating that when REG = 0, inflation amplifies the employment response to changes in output. In particular, we find that moving from 0 percent inflation to 5 percent inflation leads to a five-fold increase in the employment Okun coefficient (first row of Table 3). This finding can be interpreted as evidence of a sand effect of inflation. We also find that the coefficient attached to *DYREG* is positive and statistically significant. This indicates that labor market regulations increase the elasticity of employment to changes in output (this is the opposite of what Bertola, 1990, finds in a cross-section of nine industrialized countries). The effect is extremely large in presence of zero inflation. In this case, increasing labor market regulations from zero to 0.25 (just above the average in industrial countries) increases the Okun coefficient by approximately seven times (first column of Table 3).

As expected, we find that the coefficient attached to *DYINFREG* is negative and statistically significant. This indicates that the sand effect of inflation decreases when labor market regulations increase. In fact, when REG is equal to 0.25, inflation becomes neutral (second row of Table 3),[13] and when REG is very high (0.4 or above) inflation starts greasing the wheels of the labor markets by substantially reducing employment elasticity. In particular, when REG = 0.5 moving from 0 to 5 percent inflation reduces the employment Okun coefficient by exactly 50 percent (row three of Table 3). We take these results as evidence that inflation does grease the wheels of the labor market—but only when they are rusty. When the labor market is flexible, inflation only has a sand effect.

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and United States. For Denmark, we only have data for the 1996-1999 period. In the case of the Netherlands and Sweden we do not have data for the 1993-1996 period. For Germany, we drop 1991and 1992 (because of the unification process). For New Zealand, we only have data for the 1990-1999 period. In the case of Greece, Portugal, and Spain we drop the 1980s because there are large outliers (the basic results do not change if Greece and Spain are included in the sample but they do not hold if Portugal is included). The results are robust to using only the countries for which we have a full sample. In the sample of developing countries, the number of observations ranges from 17 (for Chile, Colombia, Costa Rica, Panama and Trinidad and Tobago) to two (Uruguay).

[13] It is exactly neutral when REG = 0.2.

Column 2 of Table 2 repeats the exercise of column 1 by restricting the sample to countries for which we do not have missing observations (so the panel of column 2 is balanced with 17 observations for each country). The results are unchanged.

The use of a fixed effect model in the presence of a variable that has limited temporal variation (like our index of labor market regulations) could be problematic because the high correlation of such a variable with country fixed effects may exacerbate measurement error and greatly increase the noise-to-signal ratio. It should be pointed out, however, that while it is true that our index of labor market regulations has limited over time variation (this is why we are not particularly interested in $b_2$), what we are interested in is the interaction between changes in output, labor market regulations, and inflation. This is a variable that does have substantial variation over time. In any case, we check for possible problems with the fixed effects specification by re-estimating the same model using a random effects specification (column 3). The results of the two models are almost identical. The only difference is that, as expected, REG is statistically significant in the random effect model but not in the fixed effect model.

Column 4 estimates the same model of column 1 by substituting employment and GDP growth with their deviations from a log-linear trend. Again, the results are unchanged. Next, we run a STATA robust regressions procedure to check whether the results are driven by outliers (column 5).[14] The results (both the magnitude of the coefficients and the t-statistics) are very similar to those of column 1, indicating that our results are not driven by outliers.

Another possible problem with the estimation of Equation 12 is the presence of the lagged dependent variable that may introduce a bias in the estimation of a fixed effect model. Column 6 addresses this issue by using the, by now standard, first difference GMM estimator originally proposed by Arellano and Bond (1991). Again, we find no major difference with respect to the coefficients and t-statistics of column 1. The last two rows of the table show that the over-identifying restrictions are valid (the Sargan test does not reject the null). While we reject the null hypothesis of no first-order correlation in the residual, we cannot reject the null of no second-order correlation (the presence of a second-order correlation would lead to inconsistent estimators).[15] Column 7 runs the Arellano and Bover (1995) system

[14] This procedure starts by eliminating all outliers for which Cook's distance is greater than one. Next, it weighs outliers by performing Huber and biweight iterations (STATA, 2002). We obtain the same results by running quantile (median) regressions.

[15] The results of the model also agree with Bond's (2002) rule of thumb for a well-specified GMM first difference model. In particular, he discusses that OLS estimates should provide an upper bound for the coefficient of the lagged dependent variable, fixed effect estimations a lower bound, and GMM estimations should be a convex combination of the two. This is exactly what we find. The point estimate of the coefficient attached to the lagged dependent variable is higher than the coefficient obtained with the fixed effect regression and lower than the one obtained with OLS (0.42, full OLS estimations not reported). The GMM estimations reported in column 6 use all the available lags of the explanatory variables as instruments. The results are robust to using different lag structures.

GMM estimator. Again, the results are unchanged. However, the high value of the Sargan test indicates that there may be problems with the specification.[16]

Next, we recognize that our results may be driven by the presence of reverse causality. It is in fact likely that a drop in employment would cause a drop in aggregate demand and hence a drop in GDP. Therefore, *DY* (and *DYINF*, *DYREG*, *DYINFREG*) is not exogenous with respect to *DE*. We address this issue by instrumenting *DY* (and *DYINF*, *DYREG*, *DYINFREG*) with an external demand shock measured by the trading partner's GDP per capita growth (weighted by trade share). This variable has all the characteristics of a good instrument, as it is highly correlated with GDP growth and it is unlikely to have a direct effect on employment (or on employment elasticity). Column 8 shows that the instrumental variable estimates yield coefficients that are essentially identical to the ones of the fixed effect regression. In this case, however, we have loss of precision. The coefficient attached to *DYINF* is no longer statistically significant. However, the coefficient attached *DYINFREG* (and to *DYREG*) remains statistically significant (although its value drop substantially and its p-value increase from 0.02 to 0.09).

Finally, we check whether our results are robust to using different indexes of labor market regulation. Columns 9 and 10 of Table 2 use the Botero et al. (2003) indices of job security and industrial relations law, while column 11 uses the Nickell et al. (2001) employment protection index.[17] The three columns yield the same message: employment elasticity is higher in highly regulated labor markets, and the effect of job elasticity decreases with inflation.

Overall, we take the results of Table 2 as providing strong evidence that in industrial countries there is a robust correlation between the interaction of labor market regulations and inflation and the Okun coefficient that measures employment elasticity with respect to output changes.[18]

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[16] All the estimations used in the paper were obtained by using STATA with the exception of the System GMM estimations that were obtained using the OX-DPD package. In the GMM-system estimations our set of instruments include 4 lags of the dependent variables. The results are unchanged if we use longer lag structure, but convergence takes much longer.

[17] One drawback of the Botero et al. indexes is that they are measured for the late 1990s and have no overtime variation. In order to use it in our panel regression, we make the assumption of no changes in labor regulations. For the employment protection index the data extend only until 1995; afterward the assumption was made of no changes in labor regulations. These regressions do not include New Zealand, which is a large outlier.

[18] There are two caveats with the results of Table 2. First of all, the results collapse if we run separate regressions for the 1980s and 1990s. This is probably due to the fact that the parameters are identified by the fact that average inflation decreased substantially from one decade to the other. (In the 1980s average inflation in OECD countries was 7.8 percent, while in the 1990s it was 2.6 percent). Second, while one would expect that the role of inflation

### *Evidence from Developing Countries*

Table 4 reproduces the same regressions of Table 2 for a sample of developing countries and Latin American countries (column 2).[19] We find that in most specifications inflation and labor market regulations do not significantly affect how employment responds to changes in output. In most cases we even find that *DYINFREG* has a positive sign (statistically significant in two cases), which is the opposite of what we expected. We take the evidence of Table 4 as indicating that there is no strong evidence that inflation and labor market regulations affect employment elasticity in developing countries.

There are four possible reasons why we do not find any significant correlation between the interaction of inflation and labor market regulations and employment elasticity in developing countries. First of all, the lack of results may be due to the fact that the explanatory variables are measured with less precision in developing countries. In this case, the lack of a statistically significant result could be purely due to attenuation bias. Second, the result may be due to the presence of widespread indexation mechanisms that completely or partly offset the grease effect of inflation (Argentina and Brazil had indexation mechanisms until the early 1990s, and Chile still has one). Third, because of lack of enforcement, labor market regulations may not be binding. In this case, *de jure* regulations would be very different from *de facto* regulations explaining the lack of a statistically significant relationship between inflation, *de jure* labor market regulations, and employment elasticity. A fourth and related explanation has to do with the presence of a large informal sector. As a result, developing countries may end up having high levels of labor market flexibility even in the presence of strict regulations (see, for instance the discussion in Calvo and Mishkin, 2003).[20]

To control for the fact that *de jure* labor market regulations may differ from *de facto* labor market regulations, we divide our sample of developing countries into two groups. The first group contains all the country-years where the ICRG index of rule of law takes a value of 4 or higher (4 is the minimum value of rule of law in our sample of industrial countries). This is the group where *de jure* regulations are likely to coincide with *de facto* regulations. The second group includes countries with low rule of law (the ICRG index takes values below 4). In this subgroup, labor market regulations are likely to be less stringent (either because they are not applied or because there is a larger informal sector) than what would

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should be particularly strong during recessions, our results are not robust to dropping periods of economic expansion.

[19] Because of data availability (especially on labor market regulations) our sample only includes four non-Latin American developing countries: Hungary, South Korea, Poland, and Turkey. We drop Turkey from the regression because it never meets the requirement of inflation below 30 percent.

[20] Yet another explanation is related to the fact that we assumed a correlation between wage and employment rigidity. However, this correlation is likely to be weaker in developing countries that are not characterized by centralized wage bargaining (we would like to thank Carmen Pagés for pointing this out).

be predicted by their *de jure* value.[21] Table 5 shows the results of a set of regressions that separate the effect of labor market regulations in countries with high and low rule of law. The first column of the table runs a fixed effects regression for the complete sample of developing countries. The second column uses a random effects model, the third column uses robust regression, and the last two columns use the Arellano and Bond and Arellano and Bover GMM estimators. We now find that the coefficient attached *to DYINFREG* is always negative and statistically significant in countries with high levels of rule of law (*DYINFREGHRL*), while we find that the coefficient is never significant and always positive for countries with low levels of rule of law (*DYINFREGLRL*). These regressions seem to suggest that inflation does grease the wheels of the labor market in developing countries with large and effective labor market regulations.[22] These results should be taken with caution, however, because they are not robust to alternative definitions of high and low rule of law.

### *Other Robustness Checks*

Before concluding this section, we run two others robustness checks (we run the robustness checks only for the sample of industrial countries). First, we test whether the results of Table 2 are robust to using unemployment instead of employment. Table 6 shows that the results are essentially identical (the dependent variable is the negative of the change in unemployment so that the coefficients have the same interpretation as the coefficients in Table 2). In fact, the unemployment regressions yield higher t-statistics. Next, instead of yearly observations, we use a panel in which the observations are averaged over six three-year periods.[23] This robustness test is important for at least two reasons. First, employment responds to changes in output with a lag. Second, the theoretical model does not have clear indications on whether we should use current or lagged inflation and averaging variables provides a useful robustness test (if instead of using the level of inflation we use its deviation with respect to a linear of HP trend, we obtain results that are similar to the ones described above). It should be pointed out that adjustment costs (which are affected by labor market regulations) are a key determinant of employment elasticity. As the longer the period of observation, the less important adjustment costs are, we expect that labor market regulations should have a smaller effect when we move from one-year to three-year averages.

The results are reported in Table 7. Again, we find that *DYINF* and *DYREG* have a positive coefficient, indicating that they increase employment elasticity. And *DYINFREG* has a negative

[21] One issue that we do not consider, but that it is likely to be important, is that the size of the informal sector may depend on how stringent labor market regulations are.

[22] The results are not robust to the use of instrumental variables. Endogeneity, however, should be less of a concern in this sample of developing countries (mostly Latin American) that are well known to be highly volatile because they are subject to large external shocks (IDB, 1995).

coefficient, indicating that the sand effect of inflation decreases when labor market regulations increase. The coefficients are statistically significant in the fixed effects, random effects, robust, and GMM regressions, but are not significant in the System GMM and IV specifications.

## 4. Does the Effect Go through Wage Flexibility?

The estimations reported so aimed at estimating the reduced form of a model linking labor market regulations and inflation to employment elasticity. The key assumption used to derive the model is that labor market regulations increase employment elasticity because they reduce wage flexibility and that inflation can, by increasing real wage elasticity, undo the effect of labor market regulations and hence reduce employment elasticity. Using this assumption, the model shows that countries with highly regulated labor markets and low inflation respond to shocks by adjusting employment (and hence have high employment Okun coefficients), while countries with a deregulated labor markets (or a regulated labor markets and some inflation) respond to shocks by adjusting real wages.

The idea that there is a trade-off between adjustment in real wages and adjustment in employment is confirmed by Figure 3, which shows a strong negative correlation between real wages and employment elasticity. Similar evidence was found by Fallon and Lucas (2002), who showed that real wage flexibility limited the drop in employment in the countries that were hit by the East Asian crisis. Their analysis also shows that inflation played a key role and that the adjustment in real wages was mostly due to an increase in prices rather than to a drop in nominal wages.

Since Section 3 provided evidence that the interaction between inflation and labor market regulations significantly affects employment elasticity, it is now interesting to look at whether the effect goes through real wage adjustment. We test this hypothesis by estimating the same model of Equation 12 but substituting employment growth with growth in real wages (*DW*). It should be pointed out that real wages are likely to be subject to a much larger measurement error and thus the estimations of this section need to be interpreted with some caution.

The model of Section 2 would predict a positive sign for the parameter attached to *DYINF*. As labor market regulations are assumed to increase real wage rigidity and as a consequence reduce the real wage to GDP elasticity, we expect the parameter attached to *DYREG* to be negative. At the same time, we expect the parameter attached to *DYINFREG* to be positive because, for a given level of labor market

[23] The three-year periods are: 1983-85, 1986-88, 1989-91, 1992-94, 1995-97, and 1998-2000.

regulations, inflation should increase real wage elasticity. Table 8 reports the estimates for the sample of industrial countries.[24]

Contrary to what we expected, the parameter attached to *DYINF* is always negative (but rarely statistically significant). As expected, we find that the coefficient attached to *DYREG* is always negative (but never statistically significant) and the coefficient attached to *DYINFREG* is always positive. The latter is statistically significant in the fixed effects, random effects, and robust regression. However, it is not significant in the GMM and IV regressions. It should also be pointed out that in the case of GMM estimations, the Sargan test does not reject the null hypothesis that the set of instruments is not valid, indicating that there may be problems with the specification used.[25] These results indicate that there is mixed evidence for the proposition that the interaction between inflation and labor market regulations is positively associated with wage flexibility.

## 5. Conclusions

This paper takes on again the issue of inflation's grease and sand effects on the labor market by studying the interaction between inflation and labor market regulations, and its effect on employment responses to changes in output. In the case of industrial countries, we find that, during periods of low inflation, labor market regulations increase the elasticity of employment to output. We also find that, in the absence of labor market regulations, inflation has a sand effect in the sense that it amplifies the employment responses to output changes. Nevertheless, our results show that this sand effect decreases when labor market regulations increase. In particular, at very high levels of labor market regulations, inflation plays an important role in reducing the employment responses to changes in output. This leads us to conclude that inflation does grease the wheels of the labor market, but only when they are rusty. The results are weaker when we focus on developing countries, and we provide some evidence that that this could be due to lack of enforcement.

The results of this paper are clearly preliminary, and they need to be corroborated by more detailed country-specific studies. If proven to be true, however, they would yield important implications for the conduct of monetary policy, and especially for the inflation policy adopted by the European Central Bank. In particular, our estimations suggest that countries with more rigid labor markets should allow for higher average inflation with respect to countries with more flexible labor markets. Hence, one

[24] In the sample of developing countries, real wages are likely to be measured with a much larger error. In fact, estimates for this sample yield no significant result.

[25] The model does not satisfy Bond's rule of thumb either. In fact, the coefficient attached to the lagged dependent variable is lower than the one obtained in the fixed effect specification and hence is not a convex combination of the OLS and fixed effect coefficients.

would expect to observe a tougher anti-inflation policy in the US (where the REG index is equal to zero) than in Euroland (where the average value for the REG index is 0.29). In reality, over the 1999-2001 period, inflation in Euroland (2 percent) has been lower than inflation in the US (2.8 percent). Our results also suggest that there may be problems linked to having a unique inflation target for countries with very diverse labor market institutions. In particular, a level of inflation that may be optimal for countries with flexible labor markets like Ireland and the Netherlands may be too low for countries with highly regulated labor markets like Italy and Spain.

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**APPENDIX**

| Variables | Definition and Source |
|---|---|
| E | Employment (millions) (Source: WEO-IMF, September 2002, variable: total employment with the exception of Barbados, Brazil, Chile, Colombia, Costa Rica, Guatemala, Jamaica, Panama, Puerto Rico, Trinidad and Tobago, Uruguay, U.S.A. and Venezuela whose source is ILO and it is completed applying the employment variation coefficient from WEO when necessary. Argentina and Mexico are from household surveys). |
| W | Real wage index (it is calculated by deflating the index by the CPI) (Source: WEO-IMF, September 2002, variable: hourly compensation, manufacturing sector. The source of Argentina and Colombia is ECLAC, Economic Surveys. For Hungary it is EIU, Annual World Tables, variable: average real wage index. Barbados, Bolivia, Brazil, Costa Rica, Cyprus, Dominican Republic, Ecuador, El Salvador, Guatemala, Honduras, Jamaica, Luxembourg, Mexico, Nicaragua, Panama, Peru, Puerto Rico, Trinidad and Tobago, Uruguay, U.S.A. and Venezuela come from ILO, manufacturing sector. Poland comes from IFS-IMF, CD-ROM, version 1.1.54, line 65. The data for Chile come from Boletines Mensuales, INE). |
| U | Unemployment rate (Source: WEO-IMF, September 2002, variable: unemployment rate with the exception of U.S.A. whose source is The Economic Report of the President. Cyprus comes from WEO-IMF and WDI-WB, CD-ROM, version 4.2, variable: unemployment, total (% of total labor force). The source for Argentina, Bolivia, Brazil, Chile, Colombia, Costa Rica, Ecuador, Guatemala, Honduras, Mexico, Nicaragua, Panama, Paraguay, Peru, Uruguay and Venezuela is ECLAC completed with unemployment from ILO). |
| Y | Gross domestic product (constant prices, billions of local currency) (Source: WEO-IMF, September 2002). |
| INF | Inflation: Constructed using the CPI of WEO-IMF, September 2002, with the exception of Argentina, Bolivia, Brazil, Chile, Colombia, Costa Rica, Ecuador, Guatemala, Mexico, Nicaragua, Panama, Paraguay, Peru, Uruguay, U.S.A. and Venezuela in which the CPI from IFS-IMF was used. |
| REG | Job security index, it is calculated by summing up indemnities for dismissal in months of pay plus advance notice in months of pay (Source: Updated version of the index compiled by Pagés 2002, and based on Heckman and Pagés, 2000 and Pagés and Montenegro, 1999). |
| RL | Rule of law, (Source: ICRG, variable: prslor) |
| GDP PARTNER | Trading partner's GDP per capita growth (%, weighted average by trade share) (Source: IMF (directions of trade), GDF-WB and WDI-WB). |

**Table 1. Summary Statistics**

| Variable | Obs. | Mean | Std. Dev. | Min | Max |
|---|---|---|---|---|---|
| | ALL COUNTRIES | | | | |
| DE | 496 | 0.015 | 0.022 | -0.074 | 0.090 |
| DW | 472 | 0.017 | 0.042 | -0.232 | 0.267 |
| DY | 496 | 0.030 | 0.028 | -0.144 | 0.121 |
| INF | 496 | 9.160 | 10.429 | -1.166 | 49.197 |
| REG | 496 | 0.322 | 0.209 | 0.000 | 1.000 |
| | INDUSTRIAL | | | | |
| DE | 289 | 0.010 | 0.019 | -0.074 | 0.090 |
| DW | 285 | 0.014 | 0.018 | -0.044 | 0.068 |
| DY | 289 | 0.028 | 0.020 | -0.065 | 0.103 |
| INF | 289 | 3.263 | 2.355 | -1.000 | 14.644 |
| REG | 289 | 0.202 | 0.127 | 0.000 | 0.649 |
| | DEVELOPING | | | | |
| DE | 207 | 0.023 | 0.023 | -0.068 | 0.072 |
| DW | 187 | 0.020 | 0.063 | -0.232 | 0.267 |
| DY | 207 | 0.033 | 0.037 | -0.144 | 0.121 |
| INF | 207 | 17.394 | 11.691 | -1.166 | 49.197 |
| REG | 207 | 0.490 | 0.185 | 0.177 | 1.000 |

(DE, DW and DY are employment, wage and GDP growth respectively)

**Table 2. Industrial Countries**

| | (1) | (2) | (3) | (4) | (5) | (6) | (7) | (8) |
|---|---|---|---|---|---|---|---|---|
| | Employment growth Fixed effects | Employment growth Fixed effects Balanced Panel | Employment growth Random effects | Deviation from trend Fixed effects | Employment growth Fixed effects Robust Regression | Employment growth Arellano and Bond GMM | Employment growth SYS-GMM | Employment growth IV regression |
| DY | 0.113 | 0.098 | 0.100 | 0.203 | 0.141 | 0.059 | 0.193 | 0.240 |
| | (0.81) | (0.64) | (0.74) | (2.26)** | (1.25) | (0.33) | (0.89) | (0.71) |
| DYINF | 0.090 | 0.090 | 0.085 | 0.063 | 0.092 | 0.094 | 0.077 | 0.103 |
| | (2.36)** | (2.50)** | (2.23)** | (2.65)*** | (2.99)*** | (1.95)* | (2.14)** | (1.36) |
| DYREG | 2.463 | 2.003 | 2.190 | 1.750 | 1.655 | 2.685 | 2.266 | 2.724 |
| | (3.37)*** | (2.47)** | (3.19)*** | (3.61)*** | (2.81)*** | (2.96)*** | (1.81)* | (1.83)* |
| DYINFREG | -0.416 | -0.359 | -0.357 | -0.243 | -0.343 | -0.463 | -0.371 | -0.498 |
| | (2.41)** | (2.05)** | (2.08)** | (2.43)** | (2.46)** | (2.14)** | (1.84)* | (1.69)* |
| INF | -0.001 | -0.002 | -0.002 | 0.002 | -0.002 | -0.001 | -0.002 | -0.001 |
| | (1.27) | (1.66)* | (2.08)** | (2.27)** | (2.20)** | (0.52) | (1.63) | (0.62) |
| REG | 0.084 | 2.297 | -0.053 | -0.137 | 0.000 | 0.277 | -0.068 | 0.137 |
| | (1.20) | (1.37) | (2.56)** | (1.62) | (0.01) | (2.39)** | (1.88)* | (1.65)* |
| INFREG | 0.008 | 0.008 | 0.009 | -0.001 | 0.007 | 0.004 | 0.009 | 0.009 |
| | (1.71)* | (1.84)* | (1.98)** | (0.50) | (1.91)* | (0.67) | (1.94)* | (1.26) |
| LDE | 0.373 | 0.44 | 0.402 | 0.402 | 0.455 | 0.418 | 0.346 | 0.369 |
| | (9.46)*** | (11.01)*** | (10.46)*** | (13.92)*** | (14.28)*** | (9.03)*** | (5.56)*** | (5.93)*** |
| YEAR | 0.000 | 0.000 | 0.000 | 0.001 | 0.000 | -0.000 | 0.000 | 0.000 |
| | (0.75) | (0.95) | (0.57) | (4.54)*** | (0.41) | (0.10) | (1.23) | (1.03) |
| Constant | -0.306 | -0.800 | -0.191 | -1.599 | -0.167 | 0.000 | -0.302 | -0.543 |
| | (0.83) | (1.52) | (0.56) | (4.45)*** | (0.56) | (0.00) | (1.20) | (1.12) |
| Observations | 289 | 221 | 289 | 277 | 289 | 265 | 289 | 254 |
| Countries | 21 | 13 | 21 | 20 | | 21 | 21 | 21 |
| R-squared | 0.67 | 0.75 | | 0.89 | 0.82 | | | |
| Sargan test | | | | | | 134 | 591 | |
| | | | | | | (0.83) | (1.00) | |
| AR(1) | | | | | | -7.49 | -3.09 | |
| | | | | | | (0.00) | (0.00) | |
| AR(2) | | | | | | -0.21 | -0.92 | |
| | | | | | | (0.84) | (0.36) | |

Absolute value of t-statistics in parentheses
* significant at 10%; ** significant at 5%; *** significant at 1%

**Table 2 (CONT): Industrial Countries, Different Indexes of Labor Market Regulations**

| | (9) | (10) | (11) |
|---|---|---|---|
| | Employment growth<br>Fixed effects<br>Job Security from Botero et al. | Employment growth<br>Fixed effects<br>Ind. Relations from Botero et al. | Employment growth<br>Fixed effects<br>Employment Protection from Nickell et al. |
| DY | 0.229 | 0.338 | 0.310 |
| | (1.99)** | (2.32)** | (2.49)** |
| DYINF | 0.052 | 0.055 | 0.047 |
| | (2.09)** | (2.03)** | (2.06)** |
| DYREG | 1.099 | 0.149 | 0.203 |
| | (2.69)*** | (1.32) | (1.83)* |
| DYINFREG | -0.177 | -0.041 | -0.041 |
| | (1.81)* | (1.97)** | (2.00)** |
| INF | -0.001 | -0.001 | -0.001 |
| | (0.93) | (0.68) | (0.90) |
| REG | | | 0.001 |
| | | | (0.11) |
| INFREG | 0.004 | 0.001 | 0.001 |
| | (1.20) | (1.14) | (1.17) |
| LDE | 0.366 | 0.374 | 0.398 |
| | (10.58)*** | (10.69)*** | (11.79)*** |
| YEAR | 0.000 | 0.000 | 0.000 |
| | (1.01) | (1.02) | (0.83) |
| Constant | -0.298 | -0.305 | -0.261 |
| | (1.04) | (1.04) | (0.84) |
| Observations | 341 | 341 | 334 |
| Countries | 20 | 20 | 19 |
| R-squared | 0.67 | 0.66 | 0.70 |

Absolute value of t-statistics in parentheses
* significant at 10%; ** significant at 5%; *** significant at 1%

**Table 3. Grease and Sand Effect at Different Levels of Labor Market Regulation**

| **REG/INF** | **0.00** | **2.50** | **5.00** | **7.50** |
|---|---|---|---|---|
| **0.00** | 0.11 | 0.34 | 0.56 | 0.79 |
| **0.25** | 0.73 | 0.69 | 0.66 | 0.62 |
| **0.50** | 1.34 | 1.05 | 0.75 | 0.46 |
| **0.75** | 1.96 | 1.41 | 0.85 | 0.30 |

**Table 4. Developing Countries**

| | (1) | (2) | (3) | (4) | (5) | (6) | (7) | (8) | (9) | (10) |
|---|---|---|---|---|---|---|---|---|---|---|
| | Employment growth Fixed effects | Employment growth Fixed effects LAC | Employment growth Random effects | Deviation from trend Fixed effects | Employment growth Fixed effects Robust Regression | Employment growth Arellano and Bond GMM | Employment growth Arellano and Bover SYS-GMM | Employment growth IV regression | Employment growth Fixed effects Job Security from Botero et al. | Employment growth Fixed effects Ind. Relations from Botero et al. |
| DY | 0.732 | 0.974 | 0.589 | 0.401 | 0.423 | 0.703 | 0.722 | 0.972 | 0.408 | 0.006 |
| | (1.83)* | (2.23)** | (1.53) | (0.86) | (1.36) | (2.02)** | (1.76)* | (0.06) | (1.36) | (0.01) |
| DYINF | -0.036 | -0.063 | -0.027 | -0.043 | 0.001 | -0.036 | -0.043 | 0.272 | -0.019 | -0.015 |
| | (1.25) | (1.95)* | (1.00) | (1.49) | (0.06) | (1.96)** | (1.49) | (0.47) | (0.91) | (0.51) |
| DYREG | -0.780 | -1.290 | -0.450 | -0.421 | -0.206 | -0.753 | -0.728 | -4.267 | -0.153 | 0.213 |
| | (0.97) | (1.46) | (0.58) | (0.45) | (0.33) | (1.04) | (0.96) | (0.10) | (0.29) | (0.70) |
| DYINFREG | 0.078 | 0.132 | 0.050 | 0.078 | 0.032 | 0.076 | 0.082 | -0.211 | 0.037 | 0.012 |
| | (1.26) | (1.92)* | (0.86) | (1.30) | (0.66) | (1.77)* | (1.40) | (0.11) | (0.91) | (0.57) |
| INF | -0.000 | -0.000 | -0.001 | -0.002 | -0.000 | 0.002 | 0.000 | -0.012 | 0.000 | 0.002 |
| | (0.20) | (0.00) | (0.96) | (1.55) | (0.38) | (2.37)** | (0.34) | (0.82) | (0.29) | (1.34) |
| REG | -0.048 | -0.048 | -0.022 | -0.017 | -0.054 | -0.053 | 0.005 | -0.015 | | |
| | (0.91) | (0.83) | (0.61) | (0.31) | (1.30) | (0.75) | (0.18) | (0.02) | | |
| INFREG | 0.000 | -0.001 | 0.002 | 0.004 | -0.002 | -0.004 | 0.000 | 0.017 | -0.000 | -0.001 |
| | (0.02) | (0.27) | (0.82) | (2.16)** | (0.65) | (1.97)** | (0.14) | (0.36) | (0.02) | (1.22) |
| LDE | -0.061 | -0.110 | 0.039 | 0.642 | 0.122 | -0.085 | -0.029 | -0.639 | -0.036 | -0.050 |
| | (0.82) | (1.35) | (0.57) | (10.05)*** | (2.10)** | (1.49) | (0.31) | (0.41) | (0.41) | (0.57) |
| YEAR | -0.001 | -0.001 | -0.002 | -0.001 | -0.000 | -0.000 | -0.001 | -0.001 | -0.001 | -0.001 |
| | (1.83)* | (1.91)* | (3.16)*** | (1.30) | (0.20) | (0.57) | (2.2)** | (0.22) | (1.23) | (1.14) |
| Constant | 2.084 | 2.321 | 3.127 | 1.783 | 0.147 | 0.000 | 2.052 | 1.293 | 1.650 | 1.457 |
| | (1.85)* | (1.93)* | (3.16)*** | (1.29) | (0.17) | (0.00) | (2.19)** | (0.23) | (1.23) | (1.14) |
| Observations | 145 | 124 | 145 | 119 | 145 | 137 | 141 | 132 | 112 | 112 |
| Countries | 14 | 11 | 14 | 12 | 14 | 14 | 13 | 14 | 12 | 12 |
| R-squared | 0.27 | 0.27 | | 0.64 | 0.61 | | | | 0.22 | 0.25 |
| Sargan test | | | | | | 167..34 | 358.30 | | | |
| | | | | | | (0.19) | (1.00) | | | |
| AR(1) | | | | | | -7.49 | -2.15 | | | |
| | | | | | | (0.00) | (0.03) | | | |
| AR(2) | | | | | | 0.21 | 0.10 | | | |
| | | | | | | (0.84) | (0.92) | | | |

Absolute value of t-statistics in parentheses

*significant at 10%; ** significant at 5%; *** significant at 1%

**Table 5. Developing Countries: DE jure versus DE facto**

| | (1) | (2) | (3) | (4) | (5) |
|---|---|---|---|---|---|
| | Employment growth Fixed effects | Employment growth Random Effects | Employment growth Fixed effects Robust Regression | Employment growth Arellano and Bond GMM | Employment growth Arellano and Bover SYS-GMM |
| DY | 0.674 | 0.667 | 0.517 | 0.529 | 0.573 |
| | (1.55) | (1.62) | (1.35) | (1.42) | (1.66)* |
| DYHRL | -0.150 | -0.874 | 0.258 | 0.774 | -0.859 |
| | (0.11) | (0.85) | (0.22) | (0.68) | (1.04) |
| DYINFHRL | 0.067 | 0.062 | 0.035 | 0.058 | 0.038 |
| | (1.05) | (1.06) | (0.63) | (1.15) | (1.33) |
| DYINFLRL | -0.039 | -0.029 | -0.021 | -0.027 | -0.028 |
| | (1.23) | (1.01) | (0.74) | (1.35) | (1.22) |
| DYREGHRL | 0.716 | 1.984 | -0.349 | -0.625 | 2.375 |
| | (0.24) | (0.90) | (0.13) | (0.25) | (1.49) |
| DYREGLRL | -0.728 | -0.560 | -0.340 | -0.377 | -0.459 |
| | (0.85) | (0.69) | (0.45) | (0.49) | (0.75) |
| DYINFREGHRL | -0.351 | -0.328 | -0.165 | -0.348 | -0.294 |
| | (1.91)* | (1.98)** | (1.03) | (2.15)** | (3.68)*** |
| DYINFREGLRL | 0.093 | 0.057 | 0.062 | 0.060 | 0.062 |
| | (1.40) | (0.95) | (1.07) | (1.27) | (1.31) |
| INF | -0.001 | -0.001 | -0.000 | 0.001 | -0.000 |
| | (0.49) | (1.15) | (0.31) | (1.76)* | (0.26) |
| REG | -0.047 | -0.025 | -0.050 | 0.043 | -0.004 |
| | (0.87) | (0.70) | (1.06) | (0.55) | (0.20) |
| INFREG | -0.000 | 0.002 | -0.002 | -0.003 | 0.000 |
| | (0.04) | (0.84) | (0.62) | (1.52) | (0.11) |
| LDE | -0.019 | 0.057 | 0.136 | -0.022 | 0.029 |
| | (0.25) | (0.80) | (1.98)* | (0.34) | (0.33) |
| YEAR | -0.001 | -0.001 | -0.001 | -0.000 | -0.001 |
| | (1.52) | (2.43)** | (1.07) | (0.07) | (2.26)** |
| Constant | 1.749 | 2.428 | 1.138 | 0.000 | 1.354 |
| | (1.54) | (2.44)** | (1.14) | (0.00) | (2.26)** |
| Observations | 141 | 141 | 141 | 133 | 137 |
| Countries | 14 | 14 | | 14 | 13 |
| R-squared | 0.34 | | 0.54 | | |
| Sargan test | | | | 132.80 | 347.10 |
| | | | | (0.867) | (1.000) |
| AR(1) | | | | -3.69 | -2.27 |
| | | | | (0.00) | (0.02) |
| AR(2) | | | | -1.26 | 0.60 |
| | | | | (0.21) | (0.55) |

Absolute value of t-statistics in parentheses
*significant at 10%; ** significant at 5%; *** significant at 1%

**Table 6. Industrial Countries, Unemployment Elasticity**

| | (1) | (2) | (3) | (4) | (5) | (6) | (7) | (8) | (9) |
|---|---|---|---|---|---|---|---|---|---|
| | Unemployment growth Fixed effects | Unemployment growth Random effects | Deviation from trend Fixed effects | Unemployment growth Fixed effects Robust Regression | Unemployment growth Arellano and Bond GMM | Unemployment growth Arellano and Bover SYS-GMM | Unemp. growth IV regression | Unemployment growth Fixed effects Job Security from Botero et al. | Unemployment growth Fixed effects Ind. Relations from Botero et al. |
| DY | 0.072 | 0.060 | -0.009 | 0.058 | 0.109 | 0.056 | 0.347 | 0.046 | 0.296 |
| | (0.90) | (0.78) | (0.15) | (0.75) | (1.15) | (0.49) | (1.59) | (0.70) | (3.50)*** |
| DYINF | 0.071 | 0.071 | -0.057 | 0.069 | 0.061 | 0.072 | 0.058 | 0.046 | 0.029 |
| | (3.24)*** | (3.28)*** | (3.48)*** | (3.30)*** | (2.41)** | (3.64)*** | (1.16) | (3.20)*** | (1.84)* |
| DYREG | 1.198 | 1.121 | -1.509 | 1.120 | 1.125 | 1.407 | 1.249 | 1.009 | 0.001 |
| | (2.86)*** | (2.87)*** | (4.54)*** | (2.79)*** | (2.39)** | (2.4)** | (1.31) | (4.32)*** | (0.02) |
| DYINFREG | -0.301 | -0.292 | 0.283 | -0.258 | -0.281 | -0.338 | -0.354 | -0.172 | -0.023 |
| | (3.03)*** | (3.01)*** | (4.11)*** | (2.71)*** | (2.51)** | (3.85)*** | (1.85)* | (3.07)*** | (1.91)* |
| INF | -0.002 | -0.002 | -0.000 | -0.002 | -0.002 | -0.002 | -0.001 | -0.001 | -0.001 |
| | (2.44)** | (3.16)*** | (0.56) | (2.59)** | (2.21)** | (3.4)*** | (0.39) | (3.13)*** | (1.34) |
| REG | -0.031 | -0.022 | 0.146 | -0.046 | -0.037 | -0.037 | 0.001 | | |
| | (0.77) | (1.83)* | (2.53)** | (1.20) | (0.61) | (2.00)** | (0.01) | | |
| INFREG | 0.006 | 0.007 | 0.001 | 0.006 | 0.008 | 0.008 | 0.006 | 0.005 | 0.000 |
| | (2.52)** | (2.77)*** | (0.32) | (2.28)** | (2.43)** | (3.41)*** | (1.34) | (2.77)*** | (1.08) |
| LDU | 0.330 | 0.351 | 0.417 | 0.268 | 0.336 | 0.340 | 0.214 | 0.330 | 0.351 |
| | (8.76)*** | (9.56)*** | (11.40)*** | (7.44)*** | (8.75)*** | (8.01)*** | (3.42)*** | (10.06)*** | (10.61)*** |
| YEAR | 0.000 | 0.000 | -0.001 | 0.000 | -0.000 | 0.000 | 0.000 | 0.000 | 0.000 |
| | (0.97) | (0.57) | (4.43)*** | (1.61) | (0.22) | (0.28) | (1.75)* | (1.11) | (0.87) |
| Constant | -0.219 | -0.118 | 1.105 | -0.353 | 0.000 | -0.061 | -0.605 | -0.197 | -0.159 |
| | (0.98) | (0.57) | (4.28)*** | (1.64) | (0.00) | (0.28) | (1.81)* | (1.15) | (0.91) |
| Observations | 278 | 278 | 267 | 278 | 254 | 278 | 243 | 329 | 329 |
| Countries | 21 | 21 | 20 | | 21 | 21 | 21 | 20 | 20 |
| R-squared | 0.68 | | 0.82 | 0.70 | | | | 0.70 | 0.69 |
| Sargan test | | | | | 174 | 598 | | | |
| | | | | | (0.09) | (1.00) | | | |
| AR(1) | | | | | -5.28 | -3.31 | | | |
| | | | | | (0.00) | (0.00) | | | |
| AR(2) | | | | | -2.04 | -2.20 | | | |
| | | | | | (0.04) | (0.03) | | | |

Absolute value of t-statistics in parentheses
*significant at 10%; ** significant at 5%; *** significant at 1%

**Table 7. Industrial Countries, 3-Year Average**

| | (1) | (2) | (3) | (4) | (5) | (6) |
|---|---|---|---|---|---|---|
| | Employment growth Fixed effects | Employment growth Random effects | Employment growth Fixed effects Robust Regression | Employment growth Arellano and Bond GMM | Employment growth Arellano and Bond SYS-GMM | Employment growth IV regression |
| DY | 0.083 | -0.356 | -0.105 | 0.018 | 0.260 | 1.129 |
| | (0.28) | (1.02) | (0.40) | (0.05) | (1.00) | (0.80) |
| DYINF | 0.245 | 0.274 | 0.235 | 0.234 | 0.172 | 0.160 |
| | (2.55)** | (2.33)** | (2.71)*** | (1.90)* | (2.01)** | (0.48) |
| DYREG | 0.448 | 0.754 | 0.676 | 0.597 | 0.296 | 0.004 |
| | (2.04)** | (3.01)*** | (3.41)*** | (2.08)** | (1.20) | (0.00) |
| DYINFREG | -0.132 | -0.153 | -0.166 | -0.159 | -0.075 | -0.097 |
| | (2.03)** | (1.94)* | (2.84)*** | (1.91)* | (1.34) | (0.40) |
| INF | -0.014 | -0.020 | -0.014 | -0.021 | -0.012 | -0.010 |
| | (1.70)* | (2.11)** | (1.85)* | (1.94)* | (1.25) | (0.40) |
| REG | -0.040 | -0.058 | -0.161 | -0.193 | -0.036 | 0.016 |
| | (0.36) | (2.72)*** | (1.61) | (1.17) | (1.48) | (0.08) |
| INFREG | 0.010 | 0.011 | 0.014 | 0.016 | 0.007 | 0.008 |
| | (1.93)* | (1.82)* | (3.06)*** | (2.52)** | (1.08) | (0.48) |
| LDE | -0.025 | 0.173 | -0.003 | 0.102 | 0.045 | -0.001 |
| | (0.38) | (2.38)** | (0.05) | (1.09) | (0.70) | (0.01) |
| YEAR | 0.003 | 0.001 | 0.002 | 0.001 | 0.002 | -0.000 |
| | (1.34) | (0.39) | (0.88) | (0.53) | (0.95) | (0.04) |
| Constant | 0.000 | 0.044 | 0.512 | | 0.003 | -0.093 |
| | (0.00) | (1.34) | (1.15) | | (0.08) | (0.37) |
| Observations | 84 | 84 | 83 | 64 | 81 | 81 |
| Countries | 20 | 20 | 20 | 17 | 17 | 20 |
| R-squared | 0.84 | 0.79 | | | | 0.64 |
| Sargan test | | | | 12.42 | 140.80 | |
| | | | | (0.19) | (0.24) | |
| AR(1) | | | | -2.09 | -2.66 | |
| | | | | (0.04) | (0.008) | |
| AR(2) | | | | -0.71 | -0.58 | |
| | | | | (0.48) | (0.56) | |

Absolute value of t-statistics in parentheses
* significant at 10%; ** significant at 5%; *** significant at 1%

## Table 8. Industrial Countries, Wage Elasticity

| | (1) | (2) | (3) | (4) | (5) | (6) | (7) | (8) | (9) |
|---|---|---|---|---|---|---|---|---|---|
| | Wage growth Fixed effects | Wage growth Random effects | Deviation from trend Fixed effects | Wage growth Fixed effects Robust Regression | Wage growth Arellano and Bond GMM | Wage growth Arellano and Bover SYS-GMM | Wage growth IV regression | Wage growth Fixed effects Job Sec. from Botero et al. | Wage growth Fixed effects Ind. Rel. from Botero et al. |
| DY | 0.227 | 0.264 | 0.202 | 0.270 | -0.047 | 0.129 | -0.032 | 0.306 | 0.252 |
| | (1.08) | (1.28) | (1.46) | (1.29) | (0.86) | (0.52) | (0.06) | (1.73)* | (1.11) |
| DYINF | -0.086 | -0.089 | -0.051 | -0.098 | -0.047 | -0.070 | -0.024 | -0.094 | -0.085 |
| | (1.46) | (1.48) | (1.26) | (1.68)* | (0.72) | (0.87) | (0.19) | (2.37)** | (1.97)* |
| DYREG | -1.299 | -1.139 | -0.293 | -1.504 | -1.211 | -0.844 | -1.279 | -1.241 | -0.206 |
| | (1.14) | (1.05) | (0.38) | (1.33) | (1.01) | (0.59) | (0.47) | (1.99)** | (1.12) |
| DYINFREG | 0.572 | 0.534 | 0.325 | 0.641 | 0.427 | 0.501 | 0.430 | 0.421 | 0.085 |
| | (1.99)** | (1.83)* | (1.73)* | (2.24)** | (1.36) | (1.22) | (0.74) | (2.80)*** | (2.33)** |
| INF | -0.000 | 0.002 | -0.003 | 0.002 | -0.003 | 0.000 | -0.002 | 0.000 | -0.001 |
| | (0.05) | (0.87) | (1.96)* | (0.83) | (1.53) | (0.04) | (0.50) | (0.08) | (1.03) |
| REG | 0.022 | 0.079 | 0.107 | 0.087 | -0.227 | 0.049 | -0.060 | | |
| | (0.21) | (2.25)** | (0.82) | (0.87) | (1.39) | (0.81) | (0.48) | | |
| INFREG | -0.017 | -0.021 | -0.009 | -0.025 | -0.008 | -0.016 | -0.014 | -0.012 | -0.001 |
| | (1.88)* | (2.42)** | (1.71)* | (2.81)*** | (0.88) | (0.95) | (0.85) | (2.52)** | (1.38) |
| LDW | 0.252 | 0.390 | 0.877 | 0.279 | 0.191 | 0.216 | 0.219 | 0.244 | 0.232 |
| | (4.87)*** | (8.18)*** | (21.93)*** | (5.42)*** | (3.49)*** | (3.84)*** | (3.69)*** | (4.86)*** | (4.59)*** |
| YEAR | -0.001 | -0.001 | -0.001 | -0.001 | -0.001 | -0.001 | -0.001 | -0.001 | -0.001 |
| | (2.54)** | (2.27)** | (3.63)*** | (2.19)** | (4.75)*** | (1.97)** | (2.06)** | (3.56)*** | (3.74)*** |
| Constant | 1.289 | 1.037 | 1.843 | 1.112 | 0.000 | 1.159 | 1.376 | 1.545 | 1.627 |
| | (2.58)** | (2.27)** | (3.60)*** | (2.23)** | (0.00) | (1.98)** | (2.13)** | (3.60)*** | (3.77)*** |
| Observations | 283 | 283 | 273 | 283 | 259 | 283 | 248 | 333 | 333 |
| Countries | 21 | 21 | 20 | | 21 | 21 | 21 | 20 | 20 |
| R-squared | 0.21 | | 0.69 | 0.44 | | | | 0.20 | 0.19 |
| Sargan Test | | | | | 195.11 | 538.90 | | | |
| | | | | | (0.01) | (1.00) | | | |
| AR(1) | | | | | -7.11 | -3.72 | | | |
| | | | | | (0.00) | (0.00) | | | |
| AR(2) | | | | | -0.72 | -0.77 | | | |
| | | | | | (0.47) | (0.44) | | | |

Absolute value of t-statistics in parentheses
* significant at 10%; ** significant at 5%; *** significant at 1%

**Figure 1. Industrial Relations and Job Security**

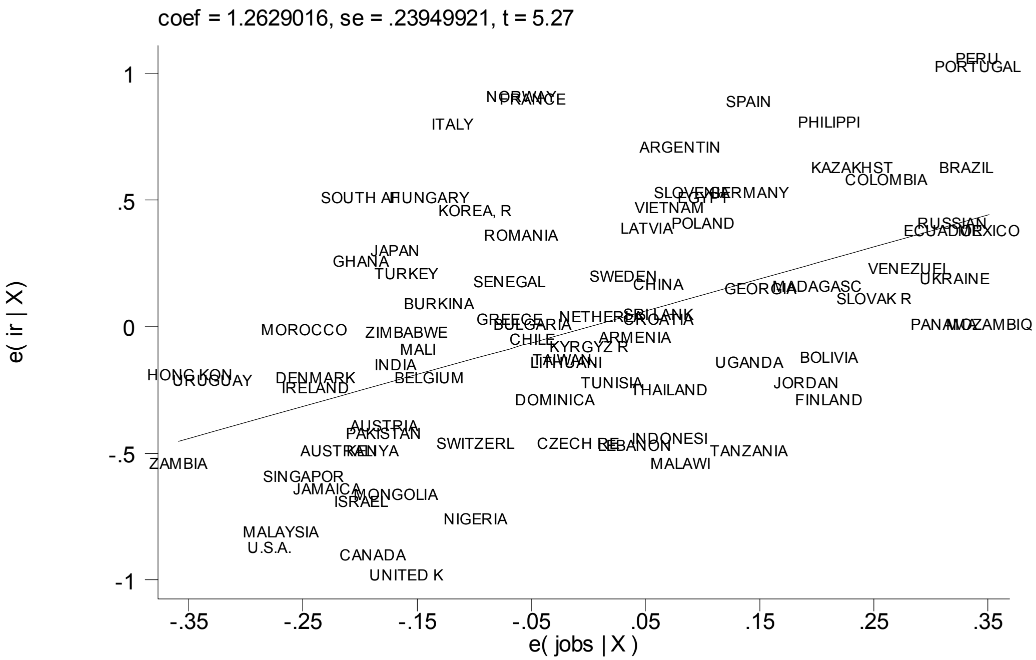

**Figure 2. Employment, Labor Market Regulations, and Inflation**

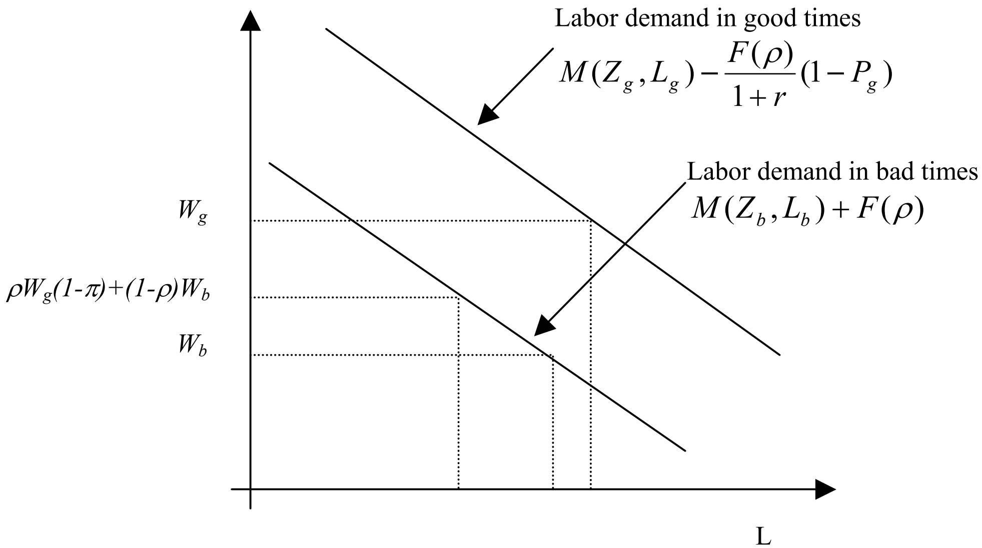

**Figure 3. Wage Elasticity versus Employment Elasticity for Industrial Countries**