Essays in Macroeconomics

by

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Abstract

This dissertation consists of three essays. The first one studies the effect of labor policy, in particular of firing costs, on financially restricted firms. It proposes and models an effect of firing costs that has not been described in the literature so far. When a time gap exists between production and its associated revenues, firing can become a liquidity adjustment tool that allows firms to increase their short-term liquidity. The presence of firing costs reduces the ability of firms to use this tool. This reduction negatively affects the optimal levels of investment and production of financially restricted firms. I present empirical evidence in line with this effect.

The second essay studies the empirical relationship between aggregate macroeconomic volatility and idiosyncratic firm-level volatility. This relationship is a testable implication of a rich set of theoretical models available in the literature. I propose a consistent estimator of the variance of firms' real sales growth rate (proxy for idiosyncratic volatility) based on the cross-sectional properties of firms' distribution. I use optimal structural break tests and long-run relationship tests to study the relationship between aggregate and idiosyncratic firm-level volatility. The main empirical results suggest a negative and significant long-run relationship.

The third essay, coauthored with Norman Loayza, analyzes potential monetary and fiscal policy biases that could result from the interaction between fiscal and monetary authorities—in a macro-policy environment where the monetary authority is committed to independently controlling inflation. We show that an increase in the divergence of authorities' preferences, with respect to the short-run trade-off between output and inflation gap, could lead to higher fiscal deficits and higher interest rates. We use a game-theoretic model to analyze this interaction, and we present supporting empirical evidence based on a panel data estimation for industrial countries.

Thesis Supervisor: Olivier J. Blanchard
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Gracias a la vida y a Candice que me han dado tanto

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Chapter 1

Labor's Liquidity Service and Firing Costs

The firing costs literature highlights two mechanisms by which firing costs affect firms' behavior: i) the efficiency cost created by the incentive to keep workers with expected marginal productivity below their wage (*e.g.*: Bentolila and Bertola 1990 and Lazear 1990); and ii) the effect that the cost of firing has on the bargaining process between employees and employers (*e.g.*: Caballero and Hammour 1998 and Blanchard 2000). This paper draws attention to an additional and independent mechanism that builds from firms' demand for liquidity. The presence of financial restrictions, and the resulting demand for liquidity, has been documented in the finance-related literature (Lamont 1997, Rauh 2005, among others). However, financial restrictions have been largely absent from the discussion of labor regulation.

The conventional views describe firing as an instrument either to adjust production to its efficiency level, or as a bargaining tool. When there is a time gap between production and its associated revenues, firing can also be understood as a liquidity adjustment tool that allows firms to increase their short-term net working capital. In other words, net liquidity is created by firing. From this perspective, a firm in need of liquidity might find it convenient to fire a worker even if his expected marginal productivity (in present value) is higher than his wage. This would allow the firm to increase its net liquidity position, and as a result, relax its financial restriction. I define this feature as *labor's liquidity service.*

The value of labor's liquidity service is affected by the presence of firing costs.

On the one hand, higher firing costs imply that more labor separations (potentially costly separations) will be needed per unit of liquidity raised by firing. On the other hand, higher firing costs could lock a firm into unavoidable situations, such as forced liquidation or the inability to invest in more profitable endeavors with minimum-scale requirements. Thus, the presence of firing costs reduces the value of labor's liquidity service, which in turn affects firms' demand for liquidity and reduces firms' demand for inputs (a downward scaling effect). I define this effect as *liquidity service effect of firing costs.*

In the case of a firm with enough internal financial resources, or with access to enough external sources of finance, an increase in firing costs will not have the above-mentioned effect. No financial restrictions are present in such a case; the demand for liquidity is zero, and therefore, the effect of the interaction between firing costs and financial restrictions does not operate.

Employment protection legislation (EPL), in the form of dismissal procedures and severance-penalty payments, is present in many countries around the world. Figure 1.1 shows the OECD's employment protection index EPL for OECD countries with available data. The ELP index shows the well known presence of high employment protection levels in some European countries *vis-a-vis* the USA. Overall, the data do not present a significant declining trend during the last decade (the average value for the late 1980's is 2.2 and for 2003 is 2.1).

The presence of employment protection, plus the existing evidence in the finance-related literature that liquidity is important for firms, suggests that the liquidity service effect of firing costs can have important macroeconomic aggregate effects. It can not only affect the medium growth rate of an economy, but also the transitional dynamics. The effects on the speed of employment recovery after a recession can be of particular interest. Some facts suggest that these effects can be important: i) small firms employment levels represent an important share of total employment and total gross job creation (*e.g.*: in the USA, small firms with 10-249 employees represent approximately 30% of total employment, and small firms with 20-99 employees represent approximately 22% of total gross job creation); and ii) small firms tend to have less, and more pro-cyclical, credit access than large firms (Gilchrist and Gertler 1994).[1]

[1]The estimates of small firms' importance are based on data from the US Bureau of Labor Statistics and from Davis *et al.* (1996).

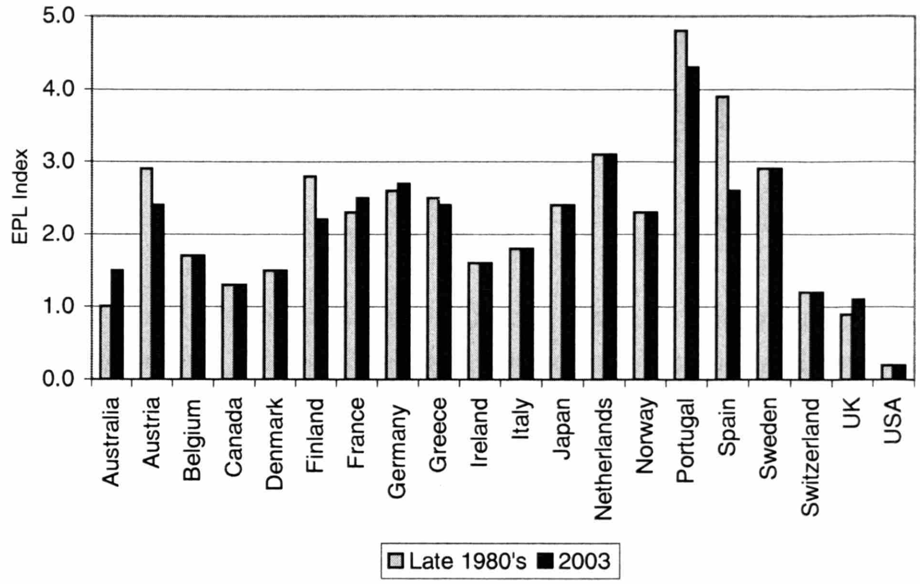


Figure 1.1. The ELP index corresponds to the OECD employment protection legislation of regular employment. Source: Employment Outlook, OECD (2004).

This paper also draws implications for studying welfare effects of firing costs. Hopenhayn and Rogers (1993), and the literature that followed [2], have characterized the welfare effect of firing costs in neoclassical general equilibrium models. In these papers, capital accumulation is only indirectly affected by firing costs—firing costs affect employment levels, and employment levels affect the marginal productivity of capital. When considering the liquidity service effect of firing costs, a new and potentially important effect arises. Operating through the firms' cash flow constraint, firing costs have a direct effect on the capital accumulation of financially restricted firms.

The paper proceeds as follows. Section 2 presents a 3-period model that studies the liquidity service effect of firing costs. Section 3 provides empirical evidence in line with the liquidity service effect of firing costs. I find a relatively stronger negative effect of firing costs on value added of industries with higher liquidity requirements. In addition, I find a relatively stronger negative effect of firing costs on value added of small firms in more labor intensive sectors. Section 4 concludes and discusses implications for labor policy design.

[2] Alvarez and Veracierto (2001) and Veracierto (2001), among others.

1.1 A Model of Labor's Liquidity Service and Firing Costs

1.1.1 The setting

Assume an entrepreneur that starts period 1 with initial wealth W and has access to a profitable investment project. The project lasts for 3 periods (investment, production, and output), uses labor and capital in fixed proportions (Leontief technology), and has constant returns to scale.

In period 1, the entrepreneur sets up a firm to undertake the project and chooses the project's scale H—the firm operates under conditions of limited liability, and the scale of the project has a maximum size of $\bar{H}$. Given the scale chosen, the firm hires H workers (*e.g.*: trainee and start-up duties), and pays eH for this investment (*e.g.*: plant and wages). The firm can fire workers in period 2 in order to keep only $L \leq H$ workers for production. The firm's labor-related cash outflow in period 2 is then $\omega L + \psi(H - L)$, where ψ represents the level of firing costs faced by the firm ($\psi \geq 0$). Note that the production technology assumes that production cannot be increased instantaneously (*e.g.*: the necessary equipment needed per worker takes one period to be installed). In period 3, the firm output is $F(L) = AL$.

The project is not free of uncertainty from the financial side. In period 2, the entrepreneur faces a stochastic liquidity shock Z that hits her level of wealth. This liquidity shock is private information for the entrepreneur—a highly valued investment or consumption opportunity (following Diamond and Rajan 2001). As a result, the disposable wealth of the entrepreneur in period 2 is $W - eH - Z$, where $Z \sim U[\underline{Z}, \bar{Z}]$.

Following Hart and Moore (1994), I assume that the outcome in period 3 is not verifiable (private information), and that firm's investment is specific to the entrepreneur. Due to the fact that the entrepreneur cannot contract upon the firm's future outcome, the access to external sources of finance is limited.

As a result of these last two assumptions, the entrepreneur has to cover the liquidity requirements of period 2 with her disposable wealth. If these liquidity needs are not financed, the entrepreneur is forced to liquidate the firm. This financial restriction creates a demand for liquidity in period 1, with the purpose of building a cash buffer stock to face the uncertainty in period 2 (insurance purposes).

1.1.2 The problem of the entrepreneur

The problem of the entrepreneur can be written as follows. Given W, the entrepreneur maximizes the value of the firm: [3]

$$V(W) = \max_{H,\ \mathbb{I}(\cdot),\ L(\cdot)} -eH + \mathbb{E}\Big(F(L) - \omega L - \psi(H-L)\Big)(1-\mathbb{I}) \tag{1.1}$$

Subject to:

$$H \leq \bar{H} \qquad L \leq H$$

$$W - eH - Z \geq \omega L + (H-L)\psi \quad \text{if } \mathbb{I} = 0$$

$$W \geq eH$$

Where the control variable $\mathbb{I}$ takes the value of 1 if the entrepreneur liquidates the firm in period 2 or the value of 0 if she does not liquidate.

Once the entrepreneur has decided the firm's scale in period 1, in period 2 she has to decide the optimal labor force L. The marginal cost of an extra unit of L is either the full cost of liquidation or $\omega - \psi$. With respect to the decision in period 1, note that the marginal cost of an additional unit of H is always higher than or equal to $e + \omega$ (higher because of the possible costs associated with the presence of financial restrictions). This is true only until the net benefit of staying in business is higher or equal to the net benefit of liquidating the firm. Since the firm is under conditions of limited liability, the entrepreneur may find it convenient to liquidate the firm if the restructuring needed is big enough.

Thus, in period 2 the optimal strategy for the entrepreneur is to hit the corner given by H or by the financial restriction, unless the necessary restructuring makes the project unprofitable. With this in mind, the optimal decision in period 2 can be written as:

$$L(W,H) = \begin{cases} H, & \text{if } Z \leq Z^F; \\ \dfrac{W - Z - H(e+\psi)}{\omega - \psi}, & \text{if } Z^F \leq Z \leq Z^L; \\ 0 \text{ (liquidation)}, & \text{if } Z^L < Z. \end{cases} \tag{1.2}$$

[3]This formulation is equivalent to assuming a risk neutral entrepreneur that maximizes the sum of her consumption from period 1 to period 3.

Where Z^F and Z^L are given by:

$$\begin{aligned} Z^F(W,H) &= \min\{W - eH - \omega H \ , \ \bar{Z}\} \\ Z^L(W,H) &= \min\{W - eH - \psi\theta H \ , \ \bar{Z}\} \ , \end{aligned} \tag{1.3}$$

and θ is defined as $\theta \equiv \frac{A}{A-\omega+\psi}$.

The function $Z^F(W,H)$ represents the maximum size of the liquidity shock, such that an entrepreneur with a firm of size H, and resources $W - eH$ at the beginning of period 2, can finance herself throughout the project. If $Z > Z^F(W,H)$ the entrepreneur does not have access to enough liquidity, unless she creates internal liquidity by firing part of the firm's labor force. Faced with this situation, the entrepreneur's optimal decision is to finance the firm by reducing its scale to L and firing $H - L$ workers.

The entrepreneur could continue financing the project only if the liquidity shock is not larger than $W - eH - \psi H$—when even firing all the work force does not create enough liquidity to finance a higher cash outflow in period 2. However, the entrepreneur will not necessarily finance the firm at that level of the liquidity shock, because she might find it convenient to liquidate the firm before reaching that point. The threshold level of the liquidity shock $Z^L(W,H)$, such that the entrepreneur decides to liquidate for any higher level, is given by the following condition: $AL - L\omega - \psi(H - L) \geq 0$. That is, staying in business is at least as profitable as liquidating the firm. Given the optimal level of labor in period 2, we can solve for $Z^L(W,H)$. [4]

The problem of the entrepreneur represented in equation 1.1 can be rewritten as:

$$V(W) = \max_{H \leq \bar{H}} -eH + \int_{\underline{Z}}^{Z^L(W,H)} F(L) - \omega L - (H - L)\psi \; dG(z) \tag{1.4}$$

Subject to the set of restrictions in equations 1.2 and 1.3 and to $W \geq eH$.

[4]Note that $\omega \geq \theta\psi$, which implies that $Z^F(W,H) \leq Z^L(W,H)$ (strong inequality unless $\psi \geq \omega$).

1.1.3 Optimal scale

The first order condition (FOC) for the problem of the entrepreneur can be expressed as:

$$0 = -e + \int_{\underline{Z}}^{Z^F(W,H)} (A-\omega)\; dG(z) + \\ \int_{Z^F(W,H)}^{Z^L(W,H)} \left[\left(A-\omega\right)\frac{-(e+\psi)}{\omega-\psi} - \psi\left(1 - \frac{-(e+\psi)}{\omega-\psi}\right)\right]\; dG(z) \tag{1.5}$$

The first line in equation 1.5 represents the standard FOC of a firm in an economy without capital market imperfections (pledgeable outcome). In this economy, enough funds would always be available as long as the project has a positive net present value ($Z^F = Z^L = \bar{Z}$). The optimal scale H in this first best world would be given by $\bar{H}$.

The second line in equation 1.5 represents the marginal effect of increasing the scale H in period 1 on the expected cost of restructuring (in the states of nature where a downward scaling is necessary to create liquidity). The first term in the integral represents the cost associated with the net marginal income loss due to the marginal increase in production destruction. The second term represents the marginal increase in firing cost expenses due to the marginal increase in destruction.

For all states of nature where the liquidity shock in period 2 is above $Z^L(W,H)$, the entrepreneur liquidates the firm and faces a continuation value of zero.

The first best allocation $\bar{H}$ is not restricted to firms in economies with perfect capital markets. In an economy with capital market imperfections, entrepreneurs with a strong financial position also choose the optimal allocation $\bar{H}$. In this model, entrepreneurs with a strong financial position are defined as entrepreneurs with initial wealth W, such that: [5]

$$W \geq \bar{W} = \bar{H}(e+\omega) + \bar{Z} - \frac{(A-\omega-e)(\omega-\psi)}{(A-\omega+\psi)(e+\omega)}(\bar{Z} - \underline{Z})$$

[5]Depending on the parameter values, the minimum W, such that $H^* = \bar{H}$, could imply that $Z^L(W) = \bar{Z}$ or that $Z^L(W) < \bar{Z}$. If $\bar{H} \geq \frac{(A-\omega-e)(\bar{Z}-\underline{Z})}{(A-\omega)(e+\omega)}$, then $Z^L(W) = \bar{Z}$. That is, if the maximum size of the project is sufficiently high, then the minimum W, such that $H^* = \bar{H}$, implies that $Z^L(W) = \bar{Z}$. Without loss of generality, and in order to simplify the presentation, I am going to assume that this condition holds.

Solving for H in equation 1.5 determines the firm's optimal scale H^* when $W \leq \bar{W}$. In deciding the optimal scale, the entrepreneur faces a trade-off between expanding the initial investment to increase profitability in the good states of nature, and taking a more conservative investment policy to reduce the cost of restructuring, or even liquidation, in the bad states of nature. The risk of the marginal unit of H decreases as initial wealth increases. As a result, the initial investment responds positively to an increase in initial wealth.

This trade-off is confirmed with the closed solution for H^*.

$$H^*(W) = \begin{cases} \bar{H}, & \text{if } \bar{W} \leq W; \\ \dfrac{W - \bar{Z} + \frac{(A-\omega-e)(\omega-\psi)}{(A-\omega+\psi)(e+\omega)}(\bar{Z} - \underline{Z})}{e+w}, & \text{if } W^B \leq W \leq \bar{W}; \\ \dfrac{W - \bar{Z} + \frac{(A-\omega-e)}{(A-\omega)}(\bar{Z} - \underline{Z})}{e + w + \frac{(e+\psi)}{(A-\omega+\psi)}\left(A - \omega + (e+\omega)\frac{\psi}{e+\psi}\right)}, & \text{if } W \leq W^B. \end{cases} \tag{1.6}$$

Where $W^B = \bar{Z} + \frac{(A-\omega-e)(Ae-\omega e+\psi e+A\psi)}{(A-\omega+\psi)(e+\omega)(A-\omega)}(\bar{Z} - \underline{Z})$ is a threshold level of initial wealth W such that below W^B it is true that $Z^L(W) < \bar{Z}$, and above W^B it is true that $Z^L(W) = \bar{Z}$. Appendix 1.A.1 shows the details of this derivation.

As expected, the optimal scale H^* increases with the level of initial wealth W (strong inequality within the interior solution range):

$$\frac{\partial H^*}{\partial W} \geq 0$$

The response of the optimal scale to an increase in initial wealth is higher at high levels of W ($W > W^B$). This is explained by the fact that there is no risk of liquidation when $W > W^B$.

This scale decision under the presence of financial restrictions parallels a portfolio problem: the entrepreneur decides how much of her wealth W to allocate in the high-return and high-risk asset (H), or in the low-return and low-risk asset (cash-liquidity). There are two additional elements in this case: i) the scale H and the marginal risk of an additional unit of H are positively related; and ii) the level of initial funds W and the marginal risk of an additional unit of H are inversely related.

1.1.4 The effect of firing costs

In this model firing costs have two effects. On the one hand, an increase in firing costs makes the restructuring more costly. For each worker in excess that the firm hires in period one (that has to be fired in the restructuring region), the firm has to fire $\frac{e+\psi}{\omega-\psi}$ additional workers. The numerator $(e+\psi)$ represents the amount of liquidity used by each worker in excess, and the denominator $(\omega-\psi)$ represents the amount of net liquidity raised by firing an additional worker.

If $\psi = 0$, the entrepreneur is fully flexible to accommodate the liquidity resources allocated to labor in period 1. In the bad states of nature, she can transform each dollar assigned to wages in period 2 into a dollar of liquidity. As ψ increases, the value of this liquidity service decreases. Creating liquidity by firing is additionally taxed and more restructuring is needed to cover the same liquidity requirements in period 2. In the limit as $\psi \to \omega$, labor hired in period 1 becomes a fully fixed production factor and $Z^L \to Z^F$.

On the other hand, an increase in firing costs decreases the region of restructuring (by increasing the region of liquidation), and increases the cost of liquidation. This last effect only applies to entrepreneurs with sufficiently low level of initial wealth $(W \leq W^B)$.

In summary, the presence of firing costs affects the firm by making the restructuring process more costly and increasing the liquidation risk. As a result, it creates the incentive to have a more conservative investment/employment policy (downward scale effect). This argument is checked by differentiating equation 1.6. [6]

$$\frac{\partial H^*}{\partial \psi} \begin{cases} < 0, & \text{if } W \leq \bar{W}; \\ \\ = 0, & \text{if } W > \bar{W}. \end{cases} \tag{1.7}$$

The optimal scale of an entrepreneur with a strong financial position $(W > \bar{W})$ is not affected by the presence of firing costs. Given her financial resources, either the risk of restructuring is zero, or small enough that the higher expected return compensates the risk at scale $\bar{H}$.

[6] To see the effect in the case that $W \leq W^B$, note that a necessary condition for the project to be profitable is that $A - \omega > e$.

The result of firing costs having no effect on unrestricted firms is not a general result. To simplify and highlight the presentation of the liquidity service effect of firing costs, the model presented in this paper shuts down the pure efficiency effect of firing costs (by eliminating the stochastic component of labor productivity). Allowing for this type of shock would not change the main conclusion: firing costs have a differentiated effect on financially restricted and unrestricted firms, creating larger distortions on financially restricted firms.

1.1.4.1 Less creation. Less destruction?

A key feature of dynamic models of firing costs is the fact that firing costs reduce the willingness to hire workers, but they also reduce the incentives to fire workers. This point is well summarized in Bentolila and Bertola's (1990) model, where they show that an increase in firing costs could have either a positive or negative effect on steady state employment level. In this sense, firing costs reduces creation, but also reduces destruction. [7] [8]

In the model presented in this paper, we already studied the financially-related negative effect that firing costs have on creation ($\frac{\partial H^*}{\partial \psi}$). With respect to destruction, there are two ways in which it could be affected by the presence of firing costs: i) affecting the effective scale of production in period 2; and ii) affecting the optimal scale in period 1 (assuming that the firm starts period 1 with a labor force greater than zero). The latter effect is equivalent to the incentive to reduce firing in Bentolila and Bertola's (1990) model.

Effective scale of production in period 2

From equation 1.7 we know that the optimal scale H^* reduces as the level of firing costs increases (for a given level of initial wealth). However, a reduction in H^* implies less risk of restructuring and/or liquidation in period 2. That is, the expected value of labor in the second period, $\mathbb{E}\, L$, could be lower because of lower H^*, or could be higher because of lower need of destruction. There is an additional effect to these dynamics: for a given level of H^*, an increase in firing costs increases the level of

[7] What is unambiguous though, is the negative effect of firing costs on turnover.

[8] In their calibrations, they find that firing costs have a small, but positive effect on steady state employment level.

destruction in the second period. This last effect goes unambiguously in the direction of lowering $\mathbb{E}\, L$.

We can use the 3-period model presented in this paper to study these dynamics. In this setting, the expected value of labor in the second period is defined as follows.

$$\mathbb{E}\, L^* = \int_{\underline{Z}}^{Z^F} H^*\, dG(z) + \int_{Z^F}^{Z^L} L(H^*)\, dG(z) \tag{1.8}$$

Note that the expected value of labor in period 2 can be written as a function of the optimal scale in period 1 and the level of firing costs: $\mathbb{E}\, L^* = \Lambda\big(H^*(\psi), \psi\big)$. Therefore,

$$\frac{\partial \mathbb{E}\, L^*}{\partial \psi} = \frac{\partial \Lambda(H^*, \psi)}{\partial H^*} \cdot \frac{\partial H^*}{\partial \psi} + \frac{\partial \Lambda(H^*, \psi)}{\partial \psi} \tag{1.9}$$

The first term in equation 1.9 represents the first two effects mentioned before. That is, the indirect effect that firing costs have over the expected production level in period 2—through the effect that firing costs have over H^*, and H^* over $\mathbb{E}\, L^*$. In the present model, the effect of reducing H^* in period 1 dominates the corresponding reduction in production destruction in period 2.

$$\frac{\partial \Lambda(H^*, \psi)}{\partial H^*} > 0 \tag{1.10}$$

Appendix 1.A.2 shows the details of this derivation. We already know that $\frac{\partial H^*}{\partial \psi} < 0$, so the indirect effect of firing costs on $\mathbb{E}\, L^*$ is also negative.

The third effect mentioned—that for a given level of H^*, an increase in firing costs increases the level of destruction in the second period—reinforces the result implied by equation 1.10. As a result,

$$\frac{\partial \mathbb{E}\, L^*}{\partial \psi} = \frac{\partial \Lambda(H^*, \psi)}{\partial H^*} \cdot \frac{\partial H^*}{\partial \psi} + \frac{\partial \Lambda(H^*, \psi)}{\partial \psi} < 0 \tag{1.11}$$

Thus, an increase in firing costs reduces the firm's initial investment/employment decision, as well as the expected level of production in period 2. In other words, the effective creation level, associated to $\mathbb{E}\, L^*$, responds negatively to an increase in firing costs.

Optimal scale and firing in period 1

How does the optimal scale H^* change if the firm starts period 1 with a labor force of size $L^- > H^*$? In the model presented in this paper, this scenario would create an incentive to reduce destruction, as in the models with no financial restrictions like Bentolila and Bertola's 1990. However, the incentive to reduce destruction is smaller than the one created in the standard models of firing costs. Even more, for certain parameter values, the fact that $L^- > H^*$ can become an incentive to *increase*, instead of reduce destruction.

When $L^- > H^*$, the problem of the entrepreneur can be written as follows. Given W and L^-, the entrepreneur maximizes the value of the firm:

$$V(W, L^-) = \max_{H,\ \mathbb{I}(\cdot),\ L(\cdot)} -eH - \psi[L^- - H]^+ + \mathbb{E}\Big(F(L) - \omega L - (H - L)\psi\Big)(1 - \mathbb{I}) \tag{1.12}$$

Subject to:

$$H \leq \bar{H} \qquad L \leq H$$

$$W - eH - \psi[L^- - H]^+ - Z \geq \omega L + (H - L)\psi \quad \text{if } \mathbb{I} = 0$$

$$W \geq eH + \psi[L^- - H]^+$$

Where the control variable $\mathbb{I}$ takes the value of 1 if the entrepreneur liquidates the firm in period 2 or the value of 0 if she does not liquidate.

The FOC of this problem is similar to the FOC of the original problem (equation 1.5), but with an additional term (the first one). For the case when $Z^L < \bar{Z}$: [9]

$$\begin{aligned} 0 = \left(\psi - \frac{\psi(A-\omega)(L^- - H)}{(\bar{Z} - \underline{Z})}\right) \mathbb{1}_{\{L^- > H\}} - e + \Big(A - \omega\Big) \int_{\underline{Z}}^{Z^F} dG(z) + \\ \int_{Z^F}^{Z^L} \left[\Big(A - \omega\Big) \frac{-(e+\psi)}{\omega - \psi} - \psi \left(1 - \frac{-(e+\psi)}{\omega - \psi}\right)\right] dG(z) \end{aligned} \tag{1.13}$$

Appendix 1.A.3 shows the details of this derivation.

[9] When $Z^L = \bar{Z}$, the additional term is $\left(\psi - \frac{\psi(A-\omega)(e+\omega)(L^- - H)}{(\omega-\psi)(\bar{Z}-\underline{Z})}\right) \mathbb{1}_{\{L^- > H\}}$. When $Z^F = \bar{Z}$, the additional term is $\psi \mathbb{1}_{\{L^- > H\}}$.

The term $\psi\mathbb{1}_{\{L->H\}}$ represents the standard incentive to reduce firing created by the presence of firing costs. If the entrepreneur has a strong financial position ($W \geq \bar{W}$), then the presence of firing costs will have the standard effect. Note that adding a positive term to the FOC implies that H^* has to increase in order to balance the FOC.

When financial restrictions are present, the term $\psi\mathbb{1}_{\{L->H\}}$ is counter-balanced with the term $-\frac{\psi(A-\omega)(L^- - H)}{(\bar{Z}-\underline{Z})}\mathbb{1}_{\{L->H\}}$, which represents the cost associated to the worsening of the firm's liquidity position. If this last term is sufficiently negative, the incentive to reduce firing would become an incentive to increase firing.

In sum, the effect behind the standard argument that firing costs reduce destruction is also present in the model shown in this paper. However, the resulting effect is the same only for financially unrestricted firms. For financially restricted firms, the fact that $L^- > H^*$ can either i) create an incentive to reduce destruction, but with less intensity than for unrestricted firms, or ii) become an incentive to increase destruction.

These results suggest that the differentiated effect of firing costs on financially restricted firms—a relatively stronger downward scaling effect compared to financially unrestricted firms—is present at the creation margin as well as at the destruction margin.

1.2 Empirical Analysis

The main conclusion from the theoretical model is that firing costs have a stronger negative effect on production levels when financial restrictions are present. I study this conclusion empirically by analyzing the effect that the interaction between firing costs and financial restrictions have on production, controlling for the corresponding individual effect of firing costs and financial restrictions. In other words, the null hypothesis tested is that the effect of firing costs does not depend on the presence of financial restrictions.

The empirical approach is based on two sets of econometric analysis. The first one centers on industry level data for a panel of countries (manufacturing sector). Borrowing from the methodology proposed by Rajan and Zingales (1998), the presence of financial restrictions is identified using cross-industry differences in the degree of liquidity requirements implied by their technology of production. The legal level of firing costs in each country, as defined by its labor law, is used as the source for layoff costs variation. Using regression analysis, I test if the effect of firing costs on industries' value added (measured as their corresponding share of national manufacturing value added) depends on their level of liquidity requirement.

One advantage of this analysis is that it allows the econometrician to control for country characteristics. This is important because country characteristics, such as financial and other labor institutional arrangements, could be correlated with the outcome measure and the legal level of firing costs. Since firing costs represent a country level characteristic, many cross-country studies that analyze the effect of firing costs do not control for this feature. [10] Another advantage of this strategy is that it relies on an exogenous source of financial restrictions variation. Centering on firms' characteristics, such as firms' levels of debt or financial ratios, could create endogeneity problems: these characteristics could determine a firm's level of access to credit markets, but they are at the same time also determined by the firm's behavior.

The empirical analysis I follow is not free of limitations. First, labor laws are not necessarily applied or supervised homogeneously across countries, which could lead to measurement errors. Second, the identification strategy regarding financial restrictions requires the assumption that the technological differences across industries are common across countries (aggregate differences in these measures across countries do not pose a problem), and that they can be computed by analyzing the behavior of firms operating in well developed financial markets (more on this later). Third, studying industry data is an indirect way of analyzing firm behavior. A more direct way would be to study firm-level data, although this approach also has its limitations. The above mentioned difficulties in identifying exogenous sources of variation in financial restrictions could be addressed using firms' size as a proxy—with the understanding that small firms tend to have limited or no access to external sources of finance. This could be done using manufacturing firm-level databases that are available for some countries. However, most of these databases use *plants* as their unit of account. *Plants* is not the relevant unit of account to address financial issues related to the plants' production level because it does not necessarily characterize the financial entity associated with that level of production. Therefore, using *plants* as

[10]This can also be the case for country-time panel data sets. As Layard (1990) states, "Unfortunately, changes in severance pay laws are quite rare, so the variation necessary to estimate within-country effects with great precision is not likely to be present in these data." (His panel consists of 22 countries and 29 years.)

the unit of account undermines the quality of the variable size as a proxy for firms' difficulty in accessing external sources of finance (as opposed to using *firms* as the unit of account).

The second set of empirical evidence studies how the importance of small firms' aggregate production—as a share of total sector production—varies with the level of legal firing costs across countries. I use firms' size as a proxy for firms' difficulty in accessing external sources of finance. This analysis is done in a difference in difference setting by comparing the importance of small firms in a relatively labor intensive sector (retail) with the importance of small firms in a relatively labor un-intensive sector (wholesale). Contrasting two sectors allows us to control for country characteristics, which, as mentioned before, is important for this type of estimations. The identification strategy assumes that the liquidity effect of firing costs is less important in the labor un-intensive sector: in the extreme case of a firm that uses no labor, the liquidity effect of firing costs is zero. The advantage of using retail and wholesale is that both are large sectors and belong to the same supply chain.

Many of the issues discussed for the first empirical analysis apply to this second analysis as well. One further challenge in this case is how to define small firms. More than one criterion can be used to identify small firms (*e.g.*: total sales, total assets, number of employees) and surely no criterion is equally applied in the case of every industry or firm. To address this issue, I use the number of employees as a criterion to identify small firms; this variable is the most standard criterion used and it is an easy variable to measure. I define a small firm as one employing fewer than 250 employees, following a large number of sources that use firms' size in order to classify firms with limited access to credit markets. Using the number of employees is also consistent with the identification strategy since many labor laws define exemptions for firms with fewer than 10-20 employees; this exemption level varies across countries.

In sum, this paper presents two econometric analyses for capturing the effect of the interaction between firing costs and financial restrictions. Although some of their features are new to the literature, these analyses are not free from some of the common econometric difficulties present in labor and corporate finance literature.

1.2.1 Empirical analysis: liquidity requirements

Let Y_{ij} be the importance of industry i's value added in country j's total manufacturing value added (measured as the corresponding ratio). The empirical model estimated is given by the following reduced-form equation:

$$Y_{ij} = \alpha_i + \alpha_j + \beta_1 FC_j \cdot LR_i + \beta_2 Z_{ij} + \eta_{ij} \tag{1.14}$$

Where FC_j measures firing costs faced by firms in country j; LR_i measures the degree of liquidity requirements implied by the particular technology of production of each industry i; Z_{ij} is a set of control variables; α_j and α_i denote country and industry aggregate effects, respectively.

The coefficient β_1 is interpreted as the differential effect that firing costs have on industries' behavior, according to their different degrees of liquidity requirements (implied by their technology of production).

1.2.1.1 Measure of liquidity requirements

I follow Rajan and Zingales (1998) methodology to identify technological characteristics of industries. They compute a financially related ratio for each US listed firm, and then construct the industry index as the median of all the observations within each industry. They propose US listed firms as a benchmark by which to measure financially related characteristics because the US financial market is the most efficient credit market in the world, and because listed firms are usually those with better access to financial markets. Therefore, the observed behavior of firms should be as close as possible to the one implied by its technology of production, as opposed to a firm's behavior conditionalized by its limited access to financial markets.

For this analysis, I use two different financially related ratios: Rajan and Zingales (1998) measure of *External Financial Dependence* (EFD for short), and Gitman (74) measure of *Cash Conversion Cycle* (CCC for short; see Raddatz (05) for a measure of CCC at industry level). The former index captures the timing and amounts of cash outflow and inflows related to investment and production, to assess liquidity requirements related to the overall investing process. The latter index captures the length in days between the cash outflow from input expenses and the cash inflow from sales, to assess liquidity requirements related to the production process.

These liquidity requirements indexes are constructed as follows: i) EFD is computed as the median of the ratio $\frac{\text{capital expenditures}-\text{cash flow from operations}}{\text{capital expenditures}}$, among all US listed firms within each industry (period 1980-89); ii) CCC is constructed in a similar fashion to EFD, and is defined as days in inventories plus days in receivables minus days in payables. Days in inventories is computed as $\frac{\text{total inventories}\cdot 365}{\text{cost of goods sold}}$, days in receivables as $\frac{\text{total receivables}\cdot 365}{\text{sales}}$, and days in payables as $\frac{\text{total payables}\cdot 365}{\text{cost of goods sold}}$ (period 1980-89). The correlation between EFD and CCC is 0.08.

1.2.1.2 Sample, data, and estimation

The sample consists of 40 countries and 28 manufacturing industries. To control for the quality of the data, all countries included have a per capita GDP higher than $1000 in 1985 (1985 dollars). See Table A1 and A2 in Appendix 1.A.4 for the composition of the sample.

The econometric approach is based on a panel data estimation with country and industry fixed effects, and robust to a heteroscedastic error structure (at country-industry level).

The dependent variable Y_{ij} is defined as industry i's share of total value added in country j's manufacturing sector. It is computed as the median observation within the period 1986-1995. Value added data is obtained from the Industrial Statistics Database (UNIDO 2005, Rev2), which provides a country-year panel dataset with disaggregated information for 28 industries across the manufacturing sector.

The source of firing costs variation is the legal levels of firing costs in each country j as defined by its labor law. Legal levels of firing costs are obtained from Botero *et al.* (03). The variable FC_j measures the level of severance and penalty payments, in weeks of pay, associated with firing a worker for economic reasons. The data presented in Botero *et al.* (03) is collected for the year 1997.

The set of variables Z_{ij} contain controls that intend to capture financial effects related to the industry variables EFD and CCC. This set includes the interaction between the level of each country's financial development with each measure of liquidity requirements (EFD and CCC); financial development is measured as domestic credit provided by the banking sector (fraction of GDP in 1985, World Development Indicators). The correlation between firing costs and financial development is -0.10.

The presence of fixed effects at both industry and country level allows us to control for general country characteristics, such as the degree of financial development, aggregate productivity, labor laws, etc., and for general industry characteristics, such as industry specific factors.

1.2.1.3 Results

Main Results

Columns 1 and 2 in Table 1 present the main empirical result of this paper. The estimated coefficient $\hat{\beta}_1$, associated with the interaction between firing costs and financial restrictions in equation 1.14, is negative and significant. This is true for both variables associated with financial restrictions, EFD and CCC.

A negative value of $\hat{\beta}_1$ indicates that an increase in firing costs has a relatively stronger negative effect in industries with higher financial requirements. Note that if financial markets worked perfectly, and as a result firms did not demand liquidity for insurance purposes, the expected value of $\hat{\beta}_1$ would be zero.

This negative coefficient can be interpreted by analyzing what happens to the threshold level $\bar{W}$ in the model presented in the theoretical section. $\bar{W}$ defines the minimum level above which the firm's financial position is strong enough so that the liquidity effect of firing costs does not distort its optimal allocation of resource ($\bar{H}$). As the liquidity requirements of production increase, the $\bar{W}$ threshold increases, too. From an industry perspective, this implies that higher firing costs would have a stronger negative effect on industries with higher levels of liquidity requirements.

The results of the core estimation also hold in a nested environment where both interactive terms are included in the regression, as well as both financial controls (Column 5 in Table 1). This estimation suggests that both measures, EFD and CCC, could be capturing different aspects related to financial restriction.

Using this last estimation, I compute the economic effect associated with the liquidity effect of firing costs. A one week reduction in firing costs is associated with a 0.86% (EFD) and 1.14% (CCC) growth differential between an industry more

TABLE 1: The Effect of the Interaction between Firing Costs and Liquidity Requirements: Main Results

Dependent Variable: Industry value added over total value added in manufacturing sector

	(1)	(2)	(3)	(4)	(5)
*Firing Costs * External Financial Dependence*	-0.113 (0.035)**		-0.111 (0.034)**		-0.099 (0.030)**
*Firing Costs * Cash Conversion Cycle*		-0.131 (0.055)**		-0.129 (0.055)**	-0.121 (0.053)**
*Financial Development * External Financial Dependence*	0.023 (0.006)**		0.020 (0.005)**	0.021 (0.005)**	0.020 (0.005)**
*Financial Development * Cash Conversion Cycle*		0.032 (0.007)**	0.031 (0.007)**	0.030 (0.007)**	0.030 (0.007)**
N	1066	1066	1066	1066	1066
R2	0.4808	0.4915	0.4945	0.4966	0.4996
Countries	40	40	40	40	40
Industries	28	28	28	28	28

Robust standard errors in parentheses: * significant at 10%, ** significant at 5%. Fixed effects control for country and industry unobserved characteristics.

affected by the liquidity service effect and an industry less affected. The distance between a more affected and a less affected industry is computed as one standard deviation in the corresponding variable used to measure liquidity requirements.

The results for the control variables in Z in all estimations are consistent with theory and with previous empirical evidence. The positive sign of these estimates indicates that industries with higher financial needs benefit more from financial development (Rajan and Zingales 1998).

In sum, the main empirical finding is the following. Controlling for country and industry characteristics, as well as financial features particular to each industry in each country, the effect of firing costs is more negative when the industry's degree of liquidity requirements is higher. Also, the estimations are consistent with previous findings related to financial effects at industry level.

Robustness Checks

This section presents alternative estimates for equation 1.14 that control for industries' labor intensity and countries' overall quality of institutions.

The level of firing costs in a country could have a differentiated effect depending on the industries' degree of labor intensity. Column 1 and 2 in Table 2 report the estimates of $\hat{\beta}_1$ controlling for the interaction between firing costs and industries' level of labor intensity. The labor intensity ranking across industries is constructed as the median value of the ratio between the number of employees in US industry i to gross fixed capital formation in US industry i (period 1986 and 1995, Industrial Statistics Database). Gross fixed capital formation is deflated by the US nonresident gross private domestic investment deflator. The correlation between labor intensity and EFD, and between labor intensity and CCC, is -0.22 and 0.37, respectively.

The negative coefficient of the interactive term between firing costs and liquidity requirements confirms the main result presented in Table 1: a relatively stronger negative effect of firing costs on industries with higher liquidity requirements. This allows us to reject the possibility that the liquidity requirement variables (EFD and CCC) could be working as proxies for the industries' labor intensity.

The coefficient related to the interaction between firing costs and labor intensity is negative and significant when using the EFD measure. This could be attributed to the negative efficiency effect of firing costs in resource allocation, in particular with respect to the labor factor. However, this result is not found when using the CCC measure. One possible explanation is the positive correlation between CCC and labor intensity.

Legal levels of firing costs could be correlated with other institutional arrangement that can affect firms with higher dependence on external sources of finance (in addition to the degree of financial development). To control for this effect, I estimate equation 1.14 adding the interactive term between liquidity requirements and a general index of countries' rule of law, where higher means more rule of law (Knack and Keefer 1995). The rule of law variable measures the citizens' willingness to accept the established institutions, to make and implement laws, and to adjudicate disputes during the year 1985. This variable also allows us to control for potential measurement error related to the difference in enforcement and supervision of labor laws across countries. The correlation between rule of law and firing costs, and between rule of law and financial development, is -0.31 and 0.25, respectively.

The third and fourth columns in Table 2 show the estimation of equation 1.14, controlling for the interaction between rule of law and liquidity requirements. The results are consistent with the results from Table 1 and Table 2, suggesting the presence

TABLE 2: The Effect of the Interaction between Firing Costs and Liquidity Requirements: Robustness Checks

Dependent Variable: Industry value added over total value added in manufacturing sector

	(1)	(2)	(3)	(4)	(5)	(6)
*Firing Costs * External Financial Dependence*	-0.124 (0.036)**		-0.101 (0.030)**	-0.076 (0.035)**		-0.065 (0.032)**
*Firing Costs * Cash Conversion Cycle*		-0.136 (0.058)**	-0.118 (0.055)**		-0.104 (0.059)*	-0.097 (0.057)*
*Firing Costs * Labor Intensity*	-0.505 (0.255)**	0.172 (0.230)	-0.071 (0.218)			
*Rule of Law * External Financial Dependence*				0.543 (0.182)**		0.51 (0.177)**
*Rule of Law * Cash Conversion Cycle*					0.405 (0.258)	0.371 (0.253)
*Financial Development * External Financial Dependence*	0.020 (0.005)**	0.021 (0.005)**	0.020 (0.005)**	0.016 (0.006)**	0.021 (0.005)**	0.016 (0.006)**
*Financial Development * Cash Conversion Cycle*	0.0310 (0.007)**	0.0300 (0.007)**	0.0300 (0.007)**	0.0310 (0.007)**	0.0270 (0.007)**	0.0270 (0.007)**
N	1066	1066	1066	1066	1066	1066
R2	0.4955	0.4968	0.4996	0.4978	0.4989	0.5048
Countries	40	40	40	40	40	40
Industries	28	28	28	28	28	28

Robust standard errors in parentheses: * significant at 10%, ** significant at 5%. Fixed effects control for country and industry unobserved characteristics.

of the liquidity service effect of firing costs. The coefficient related to the interaction between firing costs and rule of law is positive, but significant only when using the EFD measure. A positive coefficient reflects the expected positive effect of rule of law in fostering financial intermediation (*i.e.*: contracts are more easily enforced).

Asset collateralization

The previous empirical exercises concentrate on measures of liquidity requirements. An alternative way of measuring financial restrictions is to look at the firm's ability to collateralize its assets: higher collateralization means better access to external financial resources.

To test this idea I use Braun's (2003) index of collateralization or degree of tangibility of industries' assets. In a similar fashion to Rajan and Zingales (1998), he uses

TABLE 3: The Effect of the Interaction between Firing Costs and Liquidity Requirements: Asset Collateralization

Dependent Variable: Industry value added over total value added in manufacturing sector					
	(1)	(2)	(3)	(4)	(5)
*Firing Costs * Tangibility*	0.337 (0.143)**	0.339 (0.139)**	0.366 (0.159)**	0.306 (0.152)**	0.302 (0.153)**
*Firing Costs * External Financial Dependence*		-0.114 (0.034)**	-0.11 (0.033)**	-0.114 (0.034)**	-0.079 (0.035)**
*Firing Costs * Labor Intensity*			0.181 (0.265)		
*Rule of Law * Tangibility*				-0.553 (0.643)	-0.593 (0.643)
*Rule of Law * External Financial Dependence*					0.549 (0.188)**
*Financial Development * Tangibility*	-0.058 (0.018)**	-0.058 (0.017)**	-0.058 (0.017)**	-0.054 (0.018)**	-0.053 (0.018)**
*Financial Development * External Financial Dependence*		0.023 (0.005)**	0.023 (0.005)**	0.023 (0.005)**	0.019 (0.006)**
N	1066	1066	1066	1066	1066
R2	0.4838	0.495	0.4951	0.4957	0.4991
Countries	40	40	40	40	40
Industries	28	28	28	28	28

Robust standard errors in parentheses: * significant at 10%, ** significant at 5%. Fixed effects control for both, country and industry unobserved characteristics.

US listed firms to estimate the industries' availability of tangible assets as a measure of collateral. The ratio computed is $\frac{\text{net property, plant, and equipment}}{\text{total assets}}$. I will refer to this index as TAN. The correlation of TAN with EFD, CCC, and labor intensity is 0.01, -0.77, and -0.51, respectively. Given the high correlation of TAN with CCC, I exclude the latter from this set of estimations.

Table 3 shows the main empirical results of this paper replicated using the TAN index. Since higher TAN means better access to financial markets, a positive coefficient would be consistent with the liquidity service effect. The results in the first column of Table 3 show a positive estimate for the coefficient of interest, which indicate that an increase in firing costs has a relatively stronger negative effect on industries with lower levels of asset collateralization. The next columns replicate the robustness checks done for the estimations using the measures of liquidity requirement, and show that the main conclusion holds.

1.2.2 Empirical analysis: market participation of small firms

This section presents a difference in difference exercise that studies the effect of firing costs on the market participation of small firms (measured as their share of total sector value added). I use firms' size as a proxy for firms' difficulty in accessing external sources of finance, and I control for country characteristics by comparing the presence of small firms in a relatively labor intensive sector (retail) with the presence of small firms in a relatively labor un-intensive sector (wholesale).

The advantage of using retail and wholesale sectors is that they use labor and capital in significantly different proportions, that both are important sectors in the economy, and that both sectors belong to the same supply chain (both sectors exclude sale as well as maintenance and repair of motor vehicles). Using US national accounts and labor statistics, I compute the ratio of labor to capital in the retail and wholesale sectors. Labor is measured as employment and capital as stock of structure, equipment, and software. The average labor to capital ratio for retail during the period 1990-1999 is 4.4 times higher than the average ratio for wholesale. Yearly figures are appropriately deflated using the price indexes for structure and for equipment and software. [11]

The empirical analysis is based on the following formulation. Let S_{ij} be the share of small firms' value added in sector i of country j. The empirical model estimated is given by the following equation:

$$S_{ij} = \alpha_i + \alpha_j + \theta_1 FC_j \cdot D_i + \mu_{ij} \tag{1.15}$$

Where FC_j measures firing costs faced by firms in country j; D_i is a dummy variable that takes a value of one in the labor intensive sector and zero otherwise; α_j and α_i denote country and sector aggregate effects, respectively.

[11] Aggregating all 12 countries included in the sample (see next section), value added in the wholesale sector in 1999 represented 5.6% of total GDP, and value added in the retail sector represented 4.6%.

The coefficient θ_1 is interpreted as the differential effect that firing costs have on the share of small firms' value added in the more labor intensive sector *vis-a-vis* the effect on the share of small firms' value added in the less labor intensive sector.

Following the discussion in the previous section, it is important to take into consideration the levels of liquidity requirements in both sectors. If the liquidity requirements are higher in the wholesale sector, then the expected sign of θ_1 is ambiguous. On the one hand, higher firing costs could imply a heavier burden for small firms in wholesale because of higher liquidity requirements, but at the same time, a lower burden for these type of firms because of lower labor intensity. To address this issue, I study the level of EFD and CCC in both sectors using Rajan and Zingales (1998) methodology. The estimates show that wholesale has a lower EFD and a higher CCC. For both measure of liquidity requirement, the difference between the estimate for the whole sector minus the estimate for the retail sector is, however, small: -0.18 for EFD and 0.06 for CCC; these differences are expressed as shares of the max-min range observed among manufacturing industries for EFD and CCC, respectively.

1.2.2.1 Data and Results

The sample consists of 12 EU-15 countries with available sectoral data on aggregate value added for firms of different sizes, where size is defined by the number of employees working in the firm (see Table A3 in Appendix 1.A.4 for the composition of the sample). The data is obtained from Eurostat's Structural Business Statistics and covers the period 1996-2001.

Many labor laws define exemptions for firms with fewer than 10-20 employees; this exemption level varies across countries. For this reason, I define small firms as firms employing between 10 and 249 employees. Robustness checks are performed using the alternative range of 20 to 249 employees.

Figure 1.2 provides a first look at the data. It shows the difference in difference calculation, but instead of grouping countries into high and low firing costs groups, it plots the difference between the presence of small firms in retail and in wholesale ($S_{Rj} - S_{Wj}$) against the level of firing costs in each country. The data is for the year 1999, the only year with data for all 12 countries. This figure suggests that the impact of firing costs is relatively stronger in more financially restricted firms operating in a more labor intensive sector (small firms in the retail sector). Note that if financial

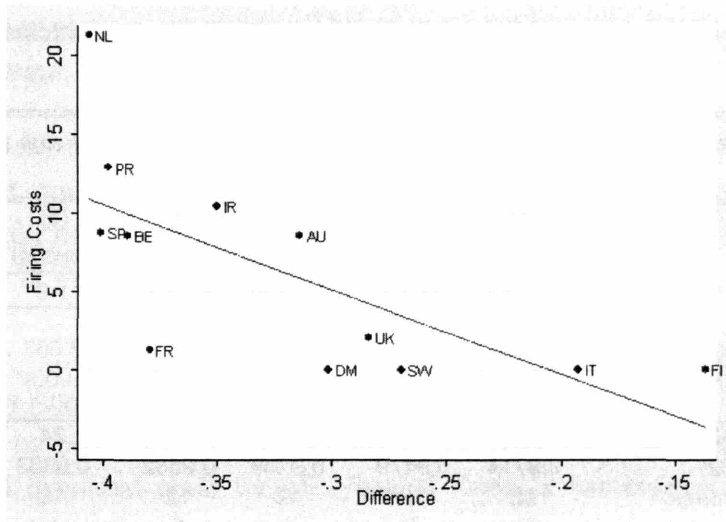


Figure 1.2. Correlation between firing costs and the difference between the small firms' share of value added in the retail and wholesale sectors.

restrictions play no role in determining the effect of firing costs, no correlation would be expected.

The negative correlation observed in Figure 1.2 is confirmed with the estimates of the parameter θ in equation 1.15. The first column in Table 4 shows the estimate of θ using only data for 1999, while the second column shows the result for the same estimation using all available data with standard errors clustered at the country level. The main conclusion from these estimates is the following: Firing costs are associated with a relatively stronger negative effect on small firms' market participation in the more labor intensive sector.

Columns 3 and 4 in Table 4 report the results for estimations equivalent to the ones presented in the first two columns of Table 4, but redefining small firms as firms employing 20 to 249 employees. The results are similar to the ones found with the alternative definition of 10 to 249 employees.

A second robustness check is performed by studying the share of value added associated with firms employing between 1 and 9 employees. Since labor laws exempt these firms from layoff costs, we should not find the negative effect of firing costs that we find in the previous estimations. These estimates are reported in the fifth and sixth columns of Table 4, showing a positive instead of negative effect of firing costs. This result is consistent with a negative effect of firing costs on financially restricted

TABLE 4: The Effect of Firing Costs on Small Firms' Market Participation: Main Results

Dependent Variable: Small firms' value added in sector *i* over total value added in sector *i* (*i*=Retail, Wholesale)						
	(1)	(2)	(3)	(4)	(5)	(6)
	Definition of a small firm (# of employees)					
	10-249	10-249	20-249	20-249	1-9	1-9
*Firing Costs * Sector Dummy (Retail=1)*	-0.009 (0.003)**	-0.008 (0.003)**	-0.006 (0.002)**	-0.005 (0.003)**	0.008 (0.003)**	0.009 (0.003)**
N	24	98	24	98	24	98
R2	0.9714	0.9479	0.9768	0.9582	0.9333	0.9308
Countries	12	12	12	12	12	12
Years	1	6	1	6	1	6

Robust standard errors in parentheses: * significant at 10%, ** significant at 5%. Fixed effects control for sector, country, and year unobserved characteristics. The standard errors are clustered at the country level for the estimates with more than one year of data.

firms, and suggests a size and/or sector selection among small firms, most likely to avoid the cost of being subject to the full extent of the labor law.

1.3 Conclusions

This work draws attention to an additional mechanism by which labor policy, in particular firing costs, affects the behavior of financially restricted firms. In short, the presence of firing costs reduces the ability to generate net liquidity through firing, and decreases the value of labor's liquidity service. As a result, there is a differentiated effect of firing costs on financially restricted firms, which is present at the creation margin as well as at the destruction margin.

Empirical evidence in line with this channel is presented. Controlling for country and industry characteristics, as well as for related financial features particular to each industry in each country, I find that the effect of firing costs on industries' value added is relatively more negative in industries with higher liquidity requirements. In addition, I find a negative effect of firing costs on the participation of small firms in more labor intensive sectors.

The main implication suggested by this paper is the following: A reduction in dismissal costs could have a proportionally higher benefit on firms for which financial restrictions are a relevant constraint. This proposition has two direct implications for labor policy design.

First, countries with less developed financial sectors, where firms face more limited access to credit, could, *ceteris paribus*, benefit more from a reduction of firing costs. This effect would be in addition to the efficiency gains associated with the standard efficiency costs of firing costs (Bentolila and Bertola 1990 and Lazear 1990).

Second, the differentiated effect of firing costs on financially restricted and unrestricted firms suggest that, *ceteris paribus*, a heterogeneous labor policy within a country could be beneficial. Compared to a homogeneous labor policy that establishes identical dismissal costs for all types of firms, a heterogeneous labor policy that establishes lower firing costs for small firms could, without changing the average level of job turnover in the economy, increase the incentive to allocate capital and labor based on productivity considerations rather than financial concerns—with the understanding that small firms are generally considered to be firms with limited or no access to credit markets (compared to large firms).

Given the importance of small firms in the gross creation of employment, this policy design could also be considered as a mechanism to promote employment creation. Many governments actively use public works, wage subsidies, and credit to small firms as a means to increase job creation, especially when unemployment is high.

In fact, labor laws in many countries do exempt smaller firms from employment protection measures. In this regard, the model presented in this paper can be understood to be a rationalization of this policy. The actual cut-off that legally defines an exempt firm changes from country to country, although in most cases it is around 10 to 20 employees. A large number of sources that use firms' size in order to classify firms with limited credit access, however, tend to define the cut-off around 250 employees. Therefore, the policy implication of this paper suggests that if financial reasons are to be taken into consideration, the legal cut-off levels used to exempt firms from paying dismissal costs might need to be reconsidered.

1.A Appendix

1.A.1 Deriving the threshold W^B

1.A.1.1 The case when $Z^L \leq \bar{Z}$

Assume a solution H^*, such that $Z^L \leq \bar{Z}$. From the first order condition in equation 1.5, we know that such a solution has to be of the following form.

$$H^* = \frac{W - \bar{Z} + \frac{(A-\omega-e)}{(A-\omega)}(\bar{Z} - \underline{Z})}{e + w + \frac{(e+\psi)}{(A-\omega+\psi)}\left(A - \omega + (e+\omega)\frac{\psi}{e+\psi}\right)} \tag{1.16}$$

To be consistent with the above statement, the following inequality has to hold.

$$W - H^*(e + \psi\theta) \leq \bar{Z} \tag{1.17}$$

Where H^* is given by equation 1.16, and $\theta \equiv \frac{A}{A-\omega+\psi}$.

Working out the inequality in equation 1.17, we can derive a condition for W, such that the solution given by equation 1.16 is consistent with the statement in equation 1.17. This condition is represented in equation 1.18.

$$W \leq \bar{Z} + \frac{(A-\omega-e)(Ae - \omega e + \psi e + A\psi)}{(A-\omega+\psi)(e+\omega)(A-\omega)}(\bar{Z} - \underline{Z}) = W^B \tag{1.18}$$

Note that a strong inequality in condition 1.17 implies that $W < W^B$.

1.A.1.2 The case when $Z^L = \bar{Z}$

Assume a solution H^*, such that $Z^L = \bar{Z}$. From the first order condition in equation 1.5, we know that such a solution has to be of the following form.

$$H^* = \frac{W - \bar{Z} + \frac{(A-\omega-e)(\omega-\psi)}{(A-\omega+\psi)(e+\omega)}(\bar{Z} - \underline{Z})}{e + w} \tag{1.19}$$

To be consistent with the above statement, the following inequality has to hold.

$$W - H^*(e + \psi\theta) \geq \bar{Z} \tag{1.20}$$

Where H^* is given by equation 1.19, and $\theta \equiv \frac{A}{A-\omega+\psi}$.

Working out the inequality in equation 1.20, we can derive a condition for W, such that the solution given by equation 1.19 is consistent with the statement in equation 1.20. This condition is represented in equation 1.21.

$$W \geq \bar{Z} + \frac{(A-\omega-e)(Ae-\omega e+\psi e+A\psi)}{(A-\omega+\psi)(e+\omega)(A-\omega)}(\bar{Z}-\underline{Z}) = W^B \tag{1.21}$$

In sum, we know from equation 1.18 that above the threshold W^B there is no solution to H^*, such that $Z^L < \bar{Z}$. Likewise, we know from equation 1.21 that below the threshold W^B there is no solution to H^*, such that $Z^L = \bar{Z}$. Therefore, it has to be true that there is only one solution for each level of W, and that such solution implies $Z^L < \bar{Z}$, for all values of $W < W^B$, and $Z^L = \bar{Z}$ for all values of $W \geq W^B$.

1.A.2 The effect of optimal scale on the expected value of labor in period 2 ($\frac{\partial \mathbb{E}\, L^*}{\partial H^*}$)

From equation 1.8, we can write the expected value of labor in period 2 as:

$$\mathbb{E}\, L^* = \begin{cases} H^* - \frac{(\bar{Z}-Z^F)^2}{2(\omega-\psi)(\bar{Z}-\underline{Z})} & \text{if } W \geq W^B \\ H^*\Big[\frac{e}{A-\omega} + H^* \frac{(e+\psi)}{(A-\omega+\psi)(\bar{Z}-\underline{Z})}\left(A-\omega+(e+\omega)\frac{\psi}{e+\psi}\right) + \\ \qquad H^*\left(\omega-\psi\theta\right)\left(1-\frac{A-\omega}{2(A-\omega+\psi)}\right)\Big] & \text{if } W < W^B \end{cases} \tag{1.22}$$

Differentiating equation 1.22:

$$\frac{\partial \mathbb{E}\, L^*}{\partial H^*} = \begin{cases} 1 - \frac{(e+\omega)(\bar{Z}-Z^F)}{(\omega-\psi)(\bar{Z}-\underline{Z})} > 0 & \text{if } W \geq W^B \\ \frac{\mathbb{E}\, L^*}{H^*} + H^* \, \frac{(e+\psi)}{(A-\omega+\psi)(\bar{Z}-\underline{Z})} \left(A - \omega + (e+\omega)\frac{\psi}{e+\psi}\right) + & \\ \qquad H^* \, (\omega - \psi\theta) \left(1 - \frac{A-\omega}{2(A-\omega+\psi)}\right) > 0 & \text{if } W < W^B \end{cases} \tag{1.23}$$

For the case when $W \geq W^B$, we know that $H^*(e+w) = W - \bar{Z} + \frac{(A-\omega-e)(\omega-\psi)}{(A-\omega+\psi)(e+\omega)}(\bar{Z} - \underline{Z})$, which implies that $H^*(e+w) - W + \bar{Z} < \frac{(\omega-\psi)}{(e+\omega)}(\bar{Z} - \underline{Z})$ (note that $\frac{(A-\omega-e)}{(A-\omega+\psi)} < 1$). This last inequality implies that $(\bar{Z} - Z^F)(e+\omega) < (\omega - \psi)(\bar{Z} - \underline{Z})$, and therefore, that $1 - \frac{(e+\omega)(\bar{Z}-Z^F)}{(\omega-\psi)(\bar{Z}-\underline{Z})} > 0$

For the case when $W < W^B$, it is straight forward to see that the derivative is positive. Just note that $A > \omega$, that $\omega \geq \psi\theta$, and that $\frac{A-\omega}{2(A-\omega+\psi)} < 1$.

In sum, an increase in the initial scale of the project (optimal) increases the production level in period 2.

1.A.3 First order condition when $L^- > 0$

Following the same steps used to derived the FOC in equation 1.5, we can write the FOC for the case when $Z^L < \bar{Z}$.

$$\begin{aligned} 0 = \psi \mathbb{1}_{\{L^- > H\}} - e + \Big(A - \omega\Big) \int_{\underline{Z}}^{W - H(e+\omega) - \psi[L^- - H]^+} dG(z) + \\ \int_{W-H(e+\omega)-\psi[L^- - H]^+}^{W-H(e+\psi\theta)-\psi[L^- - H]^+} \left[\Big(A - \omega\Big)\frac{-(e+\psi)}{\omega - \psi} - \psi\left(1 - \frac{-(e+\psi)}{\omega - \psi}\right)\right] dG(z) \end{aligned} \tag{1.24}$$

The integral $\int_{\underline{Z}}^{W-H(e+\omega)-\psi[L^- - H]^+} dG(z)$ can be written as $\int_{\underline{Z}}^{W-H(e+\omega)} dG(z) - \int_0^{\psi[L^- - H]^+} dG(z)$. Rearranging terms:

$$\begin{aligned} 0 = \left(\psi - \frac{\psi(A-\omega)(L^- - H)}{(\bar{Z} - \underline{Z})}\right) \mathbb{1}_{\{L^- > H\}} - e + \Big(A - \omega\Big) \int_{\underline{Z}}^{Z^F(W,H)} dG(z) + \\ \int_{Z^F(W,H)}^{Z^L(W,H)} \left[\Big(A - \omega\Big)\frac{-(e+\psi)}{\omega - \psi} - \psi\left(1 - \frac{-(e+\psi)}{\omega - \psi}\right)\right] dG(z) \end{aligned} \tag{1.25}$$

The FOC condition in equation 1.25 is identical to the FOC in equation 1.5 (original problem), except for the additional first term.

When $Z^L(W) = \bar{Z}$ an equivalent derivation yields the same type of analysis (see footnote 9).

1.A.4 Composition of the samples

Table A1: Manufacturing Sample (Countries)

Country	Number of Observations	Country	Number of Observations
Argentina	28	Malaysia	28
Australia	28	Mexico	26
Austria	28	Netherlands	27
Belgium	19	New Zealand	26
Brazil	18	Norway	28
Canada	28	Panama	25
Chile	28	Poland	28
Colombia	28	Portugal	28
Denmark	28	Romania	27
Ecuador	28	Singapore	26
Finland	28	South Africa	28
France	26	Spain	28
Greece	28	Sweden	28
Hungary	28	Switzerland	13
Ireland	27	Tunisia	22
Israel	28	Turkey	28
Italy	28	United Kingdom	28
Japan	28	United States	28
Jordan	28	Uruguay	28
Korea	28	Venezuela	28
		Total	1066

Table A2: Manufacturing Sample (Industries)

ISIC Code	Industry Name	Number of Observations
311	Food products	40
313	Beverages	39
314	Tobacco	39
321	Textiles	40
322	Wearing apparel, except footwear	40
323	Leather products	37
324	Footwear, except rubber or plastic	36
331	Wood products, except furniture	40
332	Furniture, except metal	38
341	Paper and products	40
342	Printing and publishing	39
351	Industrial chemicals	40
352	Other chemicals	38
353	Petroleum refineries	36
354	Misc. petroleum and coal products	32
355	Rubber products	40
356	Plastic products	38
361	Pottery, china, earthenware	39
362	Glass and products	35
369	Other non-metallic mineral products	36
371	Iron and steel	40
372	Non-ferrous metals	36
381	Fabricated metal products	37
382	Machinery, except electrical	39
383	Machinery, electric	39
384	Transport equipment	38
385	Professional & scientific equipment	37
390	Other manufactured products	38
	Total	1066

Table A3: EU Sample

Country	Number of Observations
Austria	10
Belgium	12
Denmark	6
Finland	8
France	10
Ireland	2
Italy	12
Netherlands	6
Portugal	10
Spain	6
Sweden	8
United Kingdom	8
Total	98

Chapter 2

Idiosyncratic Volatility of US Listed Firms and the Great Moderation

During the last decade, the literature has provided a rich set of theoretical models that link microeconomic volatility with aggregate volatility (Obstfeld 1994, Acemoglu 1997, Philippon 2003, Comin and Philippon 2005, and Comin and Mulani 2005). All these theories, although not for the same reasons, imply a negative correlation between aggregate and idiosyncratic volatility.

Empirically, it has been extensively documented that aggregate volatility of the US real economy has declined during recent decades (Kim and Nelson 1999, McConnell and Perez-Quiros 2000, Blanchard and Simon 2001, Stock and Watson 2003). However, there seems to be less consensus on the evolution of idiosyncratic volatility. On the one hand, Campbell *et al.* (2001), Comin and Mulani (2004), and Comin and Philippon (2005) have identified a steady increase in firm-level volatility during the last five decades, both using real data (i.e.: output) and financial data (i.e. the value of the firm). On the other hand, Davis *et al.* (2006) shows that firm volatility for the overall mass of firms in the US economy has decreased during the last three decades. [1]

The present paper revisits the empirical evidence on the evolution of idiosyncratic firm volatility, contributing in three ways. It provides a consistent estimator of firms' real sales growth rate variance (our measure of idiosyncratic volatility) based on the cross-sectional properties of firms' real sales growth rate distribution. Second, the estimator proposed in this paper distinguishes between individual firm effects and within firm effects, associating the latter with firms' idiosyncratic volatility. Third, it formally tests for the existence of a long-run relationship between aggregate and idiosyncratic volatility as well as for the possibility of a contemporaneous structural break in both series. The existing empirical literature has not formally tested nor quantified the long-run relationship between aggregate and idiosyncratic volatility.

The empirical approach proposed in this paper has two main advantages. It does not estimate the firms' idiosyncratic volatility using the time series dimension (as most of the existing empirical literature does), which implies that the resulting time series for each firms' idiosyncratic volatility is not highly autocorrelated by construction. Although the issue that concerns us in this paper is theoretically a long-term effect, the sharp and permanent decrease of aggregate volatility observed in the mid 80's—denoted in the literature as the "Great Moderation"—makes it useful to have a clearer picture of the period-by-period dynamics of firms' idiosyncratic volatility. It is reasonable to expect that if aggregate and idiosyncratic volatility are related, this discrete jump has to be accounted for by the microeconomic evidence.

Figure 2.1 stresses this point. Panel A (left) shows the 12-months quarterly growth rate of private GDP—aggregate real GDP minus government real consumption and government real gross investment—and panel B shows the standard deviation of the 12-months quarterly growth rate of private GDP within the ±10 quarters period. The vertical line in panel A is placed in the first quarter of 1984, where we can visually observe a drastic change in the volatility of private GDP growth. The two dotted vertical lines in panel B show the range of time over which we observe a decline in aggregate volatility when looking at the standard deviation constructed using the time series dimension. In panel B, the observed decline in aggregate volatility "took" 20 quarters, while in the panel A it "took" 1 quarter. [2]

[2] Statistical techniques are used in McConnell and Perez-Quiros (2000) to identify the 1984.1 break; the authors state "In this paper, we document a structural break in the volatility of U.S. GDP growth rate in the first quarter of 1984. As a mean of understanding this dramatic reduction in volatility..."

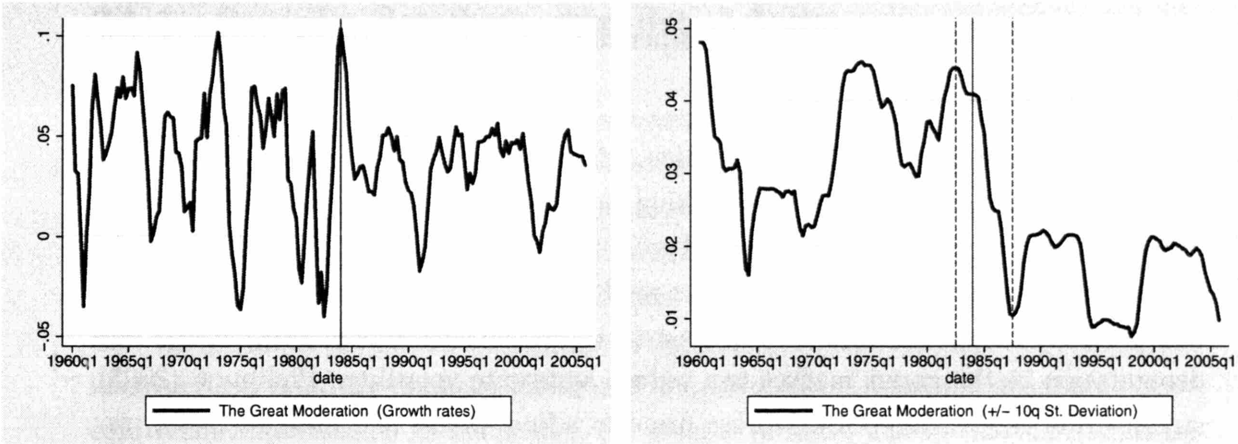


Figure 2.1.

Another advantage of the empirical approach proposed in this paper is that it estimates firms' idiosyncratic volatility free of individual firm effects. This is not only important for the cross section perspective—so as not to confuse idiosyncratic volatility with diverging trends across firms—, but also from a time series perspective—so as not to confuse idiosyncratic volatility with a firm's trend. The methodology that I use to compute individual firms effects allows for a slow moving process, which is robust to a large class of individual effect dynamics (*e.g.*: convex and concave trends as well as U-shaped and hump-shaped trends).

The paper proceeds as follows. Section 2 reviews the theoretical and empirical literature regarding the relationship between aggregate and idiosyncratic volatility. Section 3 describes the assumptions and methodology used to construct a consistent estimator of firms' real sales growth rate volatility. Section 4 presents the data and the estimate of firms' idiosyncratic volatility. Section 5 studies the long-run relationship between aggregate and idiosyncratic volatility. Section 6 presents conclusions.

2.1 Literature Review

The theoretical literature has provided a rich set of models that link firm-level microeconomic volatility with macro aggregate volatility. Two causal directions have been described in the literature. On the one hand, lower aggregate volatility can lead to higher individual risk taking, and thus, to a higher idiosyncratic firm volatility (e.g.:

macro stabilization policies). On the other hand, idiosyncratic firms' behavior can lead to lower aggregate volatility.

The channels described through which this last effect could operate are the following. First, higher individual risk taking can lead to higher diversification and as a result to lower aggregate volatility, where higher individual risk can result from an increase in financial development and/or wealth (Obstfeld 1994 and Acemoglu 1997) and/or lower macro volatility (virtuous cycle). Second, higher competition and less deregulation in the goods market can reduce aggregate volatility. Philippon (2003) stresses that more competition forces firms to adjust prices and margins faster, reducing the impact of aggregate demand shocks. According to Comin and Mulani (2005) and Comin and Philippon (2005) more competition reduces the persistence of individual firms' market shares, and as a result, increases firms' R&D investments. Higher R&D increases firm-level volatility, decreases co-movements between sectors, and reduces aggregate volatility.

These theories imply a negative long-run relationship between idiosyncratic firm volatility and aggregate volatility. Regarding the latter, it has been extensively documented that aggregate volatility of the US real economy has declined during the recent decades (Kim and Nelson 1999, McConnell and Perez-Quiros 2000, Blanchard and Simon 2001, Stock and Watson 2003); see Figure 2.1. However, there has been less consensus on the evolution of idiosyncratic volatility.

Campbell *et al.* (2001), Comin and Mulani (2004), and Comin and Philippon (2005) have identified a steady increase in listed firms' volatility during the last decades using both real data (i.e.: output) and financial data (i.e. the value of the firm). Comin and Mulani (2004) and Comin and Philippon (2005) find that the average volatility of real sales growth rate has almost doubled since the 1960's among individual US listed firms; Campbell *et al.* (2001) find that average volatility of US individual stock returns has more than doubled since the 1960's. Using a dataset with listed and non-listed firms, Davis *et al.* (2006) show that firm volatility has decreased during the last decades. In particular, they find that the average volatility of the employment growth rate for the overall mass of firms in the US economy has declined between 25% and 45%.

2.1.1 Sample: listed vs. non-listed firms

When considering the difference between listed and non-listed firms, it is important to note the dynamics that link growth, diversification, and higher individual risk taking in Obstfeld 1994 and Acemoglu 1997. As previously mentioned, these relationships are based on the idea that each agent can diversify its portfolio of risky assets/projects. The subsample of listed and non-listed firms are different with respect to the extent that these assets/projects are available to agents for diversification. Therefore, the idiosyncratic volatility found in listed and non-listed firms does not necessarily have equivalent implications. In particular, the sample of listed firms is more relevant for the class of theories that relate growth, diversification, and higher individual risk taking.

The following example emphasizes this point. Assume that agents in two economies, A and B, invest the same aggregate amount in the same number of assets/projects. However, in economy A this is done in a diversified way—each agent holds a share of each asset/project—and in economy B this is done in an individual way—each agent holds 100% of only one asset. The benefit of idiosyncratic diversification, and its positive effect on growth and in agents' willingness to take risk, operates in economy A but not in B. Agents in economy B are not diversifying their risk and therefore, are making more precautionary (less risky) investment decisions. [3]

The difference between economy A and economy B can be considered parallel to the difference between assets/projects listed in the stock market and assets/projects not listed. The ownership of the latter group is more concentrated in only a few agents' portfolios and are not available, to the same extent as listed assets/projects are, as part of the diversifiable pool of assets/projects of most agents in the economy.

Therefore, two important considerations have to be taken into account when analyzing the empirical implications of theories like Obstfeld 1994 and Acemoglu 1997—in particular when comparing aggregate volatility with idiosyncratic volatility. First, the increase in idiosyncratic volatility does not necessarily mean that each firm's volatility has to increase, but that the available pool of diversifiable assets/projects contains more risky assets/projects. The latter implication can be better tested with the sample of listed firms.

[3] Assume that agents in economy B have more wealth than agents in economy A. This helps to reconcile the fact that agents in both economies are investing identically, but agents in economy B are making more precautionary investment decisions.

Second, shouldn't the riskiness of non-listed firms increase if the riskiness of listed firms increases? On the one hand, it could be argued that if agents are more diversified in their pool of listed assets/projects, then they would be willing to take more risk in their non-listed, and individually owned, assets/projects. On the other hand, there could be a "self-selection" of assets/projects according to their implicit risk. As more agents in the economy start demanding more risky assets, they are willing to pay a relatively higher premium to access this type of assets and a relatively lower premium to access the less risky assets. This could imply that as the economy develops and grows, more risky assets are relatively more demanded for diversification purposes than less risky assets, pushing more risky assets into the diversifiable pool (stock market) and less risky assets into the non-diversifiable pool (non-listed). If the latter effect were stronger, then the evidence of the non-listed firms would give a misleading picture.

2.2 A Cross-sectional View

Campbell *et.al.* (2001), Comin and Mulani (2004), and Comin and Philippon (2005) center their empirical analyses of firms' idiosyncratic volatility on the time series variation of individual firms. They compute their preferred measure of volatility per firm (across time) and then extract a center tendency per period, either by computing the mean or the median observation among firms within each period.

In this paper, I propose an alternative empirical strategy based on two observations. First, an increase in firms' volatility should not only lead to an increase in the time series variance of each firm, but it should also lead to an increase in the cross-sectional variance among firms in a given period (unless all observations in the cross-section are perfectly correlated).

Second, if we assume that within each cross-section, firms' idiosyncratic real sales growth is independently distributed across firms, then the sample variance estimator for the cross-section sample is a consistent estimator of firms' real sales growth rate variance (our measure of idiosyncratic volatility)—see section 2.2.2.2 for the details.

The sample of firms used in this paper does not necessarily satisfy the independence assumption. To deal with this issue, I construct a measure of real sales growth rate that is independent and idiosyncratic to each firm—adjusting for aggregate and sectoral effects, in order to eliminate dependence between the cross-sectional observations.

In what follows, I describe the assumption and methodology used to construct a consistent estimator of firms' real sales growth rate variance.

2.2.1 Idiosyncratic volatility

2.2.1.1 Firm level

Let's define the stochastic process $g_{i,s}(t)$ as the family of random variables $g_{i,s,t}$ that represent the real sales growth rate of firm i in sector s in period t. The index i has a unique value for each firm and the index s only describes the corresponding sector (is redundant for indexing each individual firm).

We can decompose each random variable into 4 components:

$$g_{i,s,t} = g_t + g_{s,t} + g_i^{\mathrm{IE}} + g_{i,t}^{\mathrm{ID}} \tag{2.1}$$

Where g_t is an aggregate component common to all firms in period t, $g_{s,t}$ is a sector-period specific component common to all firms in sector s in period t, g_i^{IE} is an individual firm effect specific to each firm, and $g_{i,t}^{\mathrm{ID}}$ is a time varying and firm specific component.

The object we are interested in is the volatility of the idiosyncratic component $g_{i,t}^{\mathrm{ID}}$. In particular, we are interested in its variance and how this moment changes through time.

Let's assume that the idiosyncratic component of the real sales growth rate has an unconditional mean of zero, and that both the idiosyncratic and the individual components are orthogonal to the aggregate and sector specific components. That is,

$$\mathbb{E}\left[g_{i,t}^{\mathrm{ID}}\right] = 0 \tag{2.2}$$

$$\mathbb{E}\left[(g_t + g_{s,t}) \cdot g_{i,t}^{\mathrm{ID}}\right] = \mathbb{E}\left[(g_t + g_{s,t}) \cdot g_i^{\mathrm{IE}}\right] = 0 \tag{2.3}$$

The unconditional variance of the idiosyncratic component is $\mathbb{E}\,[g_{i,t}^{\mathrm{ID}}]^2$, where the expectations operator is over the possible realizations of the stochastic process.

For long periods of time, the assumption that g_i^{IE} is time invariant might not be a realistic assumption (e.g.: under the presence of a deterministic trend for a significant period of time). To capture possible changes in individual components, I allow for slow moving changes in g_i^{IE}. In particular, I arbitrarily model this slow moving component as follows:

$$g_{i,t}^{\mathrm{IE}} = \frac{1}{2w+1} \sum_{r=t-w}^{t+w} (g_{i,s,r} - g_{r,s} - g_r) \tag{2.4}$$

Where w is a constant and $w > 0$

2.2.1.2 Cross-section level

Let's assume that each cross-section is populated by a large number of firms n_t, with a large number of firms $n_{s,t}$ in each sector s, and that the idiosyncratic component within the cross-section is independent for each firm. In addition, let's assume that the unconditional variance of the idiosyncratic component is common to all firms within the cross-section. This last assumption implies:

$$\mathbb{E}\,[g_{i,t}^{\mathrm{ID}}]^2 = \sigma_t^{2\ \mathrm{ID}} \tag{2.5}$$

2.2.2 An estimator for $\sigma_t^{2\ \mathrm{ID}}$

2.2.2.1 Estimator $\hat{\sigma}_t^{2\ \mathrm{ID}}$

Let's define the estimator $\hat{\sigma}_t^{2\ \mathrm{ID}}$ as follows.

$$\hat{\sigma}_t^{2\ \mathrm{ID}} = \frac{1}{N_t} \sum_{\forall i} [\hat{g}_{i,t}^{\mathrm{ID}}]^2 \tag{2.6}$$

Where $N_t \leq n_t$ is a sample from the total number of firms n_t, $\forall i$ refer for all firms in the sample N_t, and

$$\hat{g}_{i,t}^{\text{ID}} = (g_{i,s,t} - \hat{g}_t - \hat{g}_{s,t} - \hat{g}_{i,t}^{\text{IE}}) \tag{2.7}$$

and

$$\hat{g}_t + \hat{g}_{s,t} = \frac{1}{N_{s,t}} \sum_{i=1}^{N_{s,t}} g_{i,s,t} \tag{2.8}$$

where $N_{s,t}$ is the number of firms in sector s for each cross-sectional sample t, and

$$\hat{g}_{i,t}^{\text{IE}} = \frac{1}{2w+1} \sum_{r=t-w}^{t+w} (g_{i,s,r} - \hat{g}_r - \hat{g}_{s,r}) \tag{2.9}$$

Where w is a constant and $w > 0$.

2.2.2.2 Consistency of $\hat{\sigma}_t^{2\ \text{ID}}$

To study the consistency of the estimator $\hat{\sigma}_t^{2\ \text{ID}}$, let's start by replacing equation (2.1) into equation (2.7) and rewriting $\hat{g}_{i,s,t}^{\text{ID}}$ as follows:

$$\hat{g}_{i,s,t}^{\text{ID}} = g_{i,s,t}^{\text{ID}} + \left(g_{i,s,t}^{\text{NID}} - \hat{g}_{i,s,t}^{\text{NID}}\right) \tag{2.10}$$

Where $g_{i,s,t}^{\text{NID}} = g_t + g_{s,t} + g_{i,t}^{\text{IE}}$ and $\hat{g}_{i,s,t}^{\text{NID}} = \hat{g}_t + \hat{g}_{s,t} + \hat{g}_{i,t}^{\text{IE}}$.

It follows that,

$$[\hat{g}_{i,s,t}^{\text{ID}}]^2 = [g_{i,s,t}^{\text{ID}}]^2 + 2[g_{i,s,t}^{\text{ID}}]\left(g_{i,s,t}^{\text{NID}} - \hat{g}_{i,s,t}^{\text{NID}}\right) + \left(g_{i,s,t}^{\text{NID}} - \hat{g}_{i,s,t}^{\text{NID}}\right)^2 \tag{2.11}$$

and that,

$$\hat{\sigma}_t^{2\ \text{ID}} = \frac{1}{N_t} \sum_{\forall i} [g_{i,s,t}^{\text{ID}}]^2 + 2\frac{1}{N_t} \sum_{\forall i} [g_{i,s,t}^{\text{ID}}]\left(g_{i,s,t}^{\text{NID}} - \hat{g}_{i,s,t}^{\text{NID}}\right) + \frac{1}{N_t} \sum_{\forall i} \left(g_{i,s,t}^{\text{NID}} - \hat{g}_{i,s,t}^{\text{NID}}\right)^2 \tag{2.12}$$

As detailed in Appendix 2.A.1, we can show that the first term in (2.12) converges to $\sigma_t^{2\ \text{ID}}$ as $N_t \to \infty$, and that the last two terms converge to zero as $N_t \to \infty$. That is,

$$\lim_{N_t \to \infty} \hat{\sigma}_t^{2\ \mathrm{ID}} = \sigma_t^{\mathrm{ID}} \qquad \left[\hat{\sigma}_t^{2\ \mathrm{ID}} \xrightarrow{p} \sigma_t^{\mathrm{ID}}\right] \tag{2.13}$$

2.3 Idiosyncratic Volatility of US Listed Firms

2.3.1 Data and sample

I use the Compustat Database sample that contains annual balance sheet data for US listed firms from 1950 to 2004. The annual real sales growth rate is constructed using the annual sales reported in each firm's income statement (deflated using US Producer Price Index).

The full sample contains 229,384 valid observations for real sales growth rate, for an average of 5,097 observations per year (576 in 1951 and 5,445 in 2004). This sample includes only positive observations of sales for firms located in the US. To avoid changes in sales due to mergers, the sample excludes year-firm sales observations affected by mergers and acquisitions. To reduce the influence of outliers, which could be due to extreme observations or measurement error, I Winsorize the first and ninety-ninth percentile—computed from the whole sample of non-international firms, which implies that only 1.39% of valid observations are removed. Finally, and to avoid a downward bias on the estimation of firms' idiosyncratic volatility $g_{i,s,t}^{\mathrm{ID}}$, observations in year-sectors with less than 6 observations are removed from the sample—they corresponds to 0.54% of valid observations. Sectors are defined based on the 2-digit SIC classification.

I use a window of 7 years to compute the individual firm effect $\hat{g}_{i,s,t}^{\mathrm{IE}}$ (w=3 in equation 2.9). Since the sample is an unbalanced panel, there is no information on the corresponding 7 years for all year-firm observations. For the computation of $\hat{g}_{i,s,t}^{\mathrm{ID}}$ I use only year-firm observations with no less than 4 out of 7 years of information. This implies that 7.65% of valid observations are not used in the estimation of $g_{i,s,t}^{\mathrm{ID}}$—one out of five correspond to observations within the period 2002-4. Table A1 in the Appendix 2.A.2 reports the sample composition per year used in the computations.

2.3.2 Estimator $\hat{g}^{\mathrm{ID}}_{i,s,t}$

Panel A (left) of Figure 2.2 plots the time series for the estimator of firms' idiosyncratic volatility $\hat{g}^{\mathrm{ID}}_{i,s,t}$.

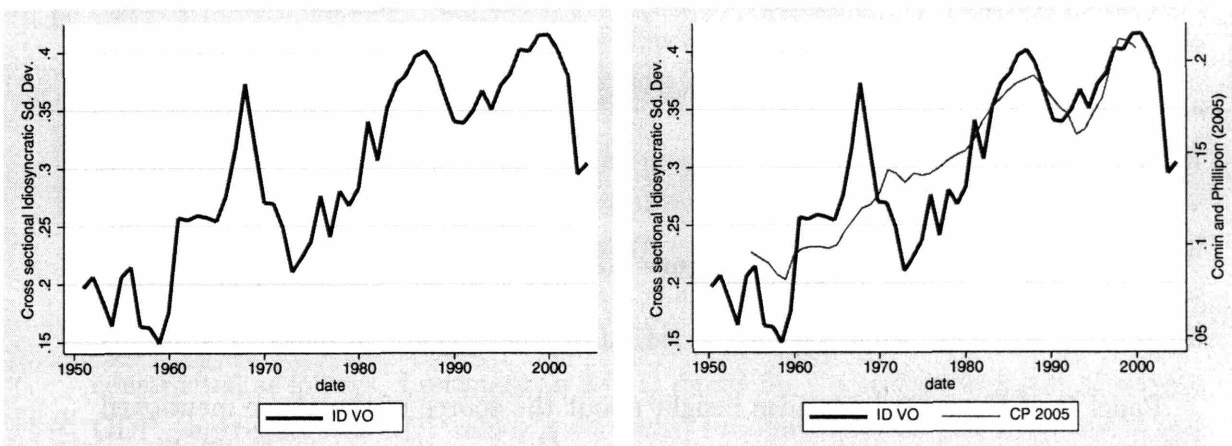


Figure 2.2.

The estimate of firms' idiosyncratic volatility presented in this paper is consistent with the empirical literature that finds a positive trend in firms' idiosyncratic volatility. In particular, panel B (right) of Figure 2.2 analyzes the parallels between the series constructed in this paper and the one reported in Comin and Philippon (2005). Notice however, that the scale of both series is significantly different. Since I argue that $\hat{g}^{\mathrm{ID}}_{i,s,t}$ is a consistent estimator of firms' idiosyncratic volatility, it is worthwhile exploring the differences between these two approaches. The difference between the cross sectional view and the time series view can also be noted in the graphs presented in Davis *et al.* (2006).

We should point out the sample used is not behind the difference reported. Panel A of Figure 2.3 shows the estimate of Comin and Philippon's (2005) estimator using the sample used in this paper to compute $\hat{g}^{\mathrm{ID}}_{i,s,t}$. The original series reported in Comin and Philippon (2005) and the one estimated with the sample used in this paper are consistent.

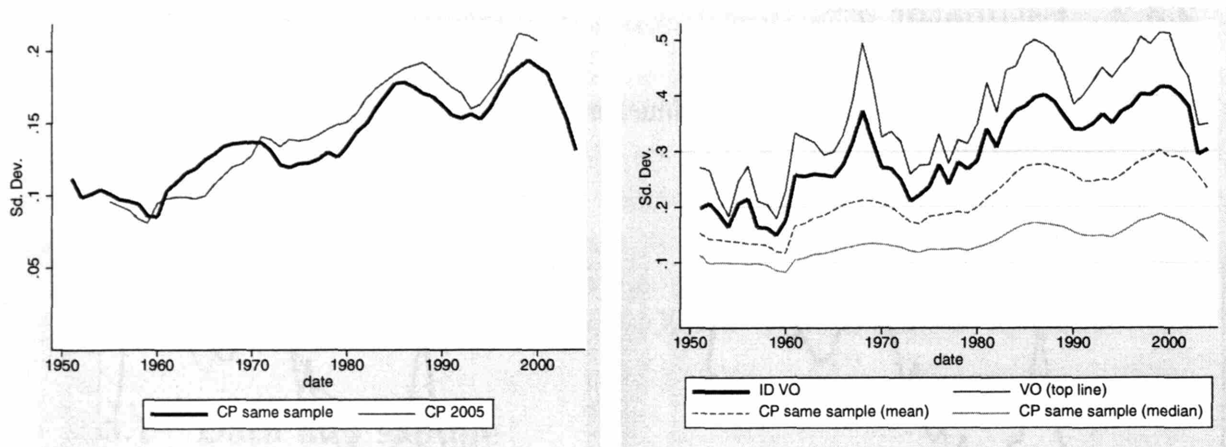


Figure 2.3.

Panel B of Figure 2.3 provides insight about the source of the above mentioned difference. The top line is the series constructed based on the cross-sectional dispersions without any aggregate, sectoral, nor individual-effect adjustment (equivalent to the cross-sectional estimate in Davis *et al.* 2006). The bottom line is the series constructed by Comin and Philippon (2005). On average, the former estimate is 169% higher that the latter. The second series from top to bottom is the estimate of $\hat{g}^{\mathrm{ID}}_{i,t}$—which accounts for aggregate, sectoral, and individual-effect adjustment in order to eliminate dependence between the cross-sectional observations—and the second series from bottom to top is the estimator used in Comin and Philippon (2005), but computing the mean observation within each cross-section instead of the median observation as Comin and Philippon (2005) do—computing the mean-version is more consistent with the estimator $\hat{g}^{\mathrm{ID}}_{i,t}$.

Although the gap is not closed completely when comparing the adjusted cross-sectional estimate with the mean version of Comin and Philippon (2005), the unaccounted difference is now 39%. This result makes us more confident on the degree of consistency of $\hat{g}^{\mathrm{ID}}_{i,s,t}$, and as the graphs show, on its changes through time. [4]

For the rest of the paper I will concentrate the analysis on the adjusted cross-sectional estimator of firms' idiosyncratic volatility (*i.e.*: $\hat{g}^{\mathrm{ID}}_{i,s,t}$).

[4] It is still worthwhile to continue to explore the sources of this 39% gap, which can help us to have a more accurate estimate of firms' idiosyncratic volatility.

2.4 Long-run Relationship between Idiosyncratic and Aggregate Volatility

As detailed in section 2.1 there is a rich set of theories that suggest a negative relationship between microeconomic volatility and macro aggregate volatility. The empirical literature has addressed this issue by constructing measures of idiosyncratic volatility, but has not formally tested the implied relationship with aggregate volatility.

In this section, I formally test for the present of a long-run relationship between aggregate volatility and idiosyncratic volatility during the period 1951-2004. I borrow from McConnell and Perez-Quiros' (2000) methodology to compute a measure of aggregate volatility that does not depend on a rolling window and I use the estimator $\hat{g}^{\text{ID}}_{i,s,t}$ as a measure of idiosyncratic volatility. The aggregate volatility measure is constructed as follows. I estimate an AR(1) model for the real growth rate of private GDP—aggregate real GDP minus government real consumption and government real gross investment. I compute the aggregate volatility measure as the absolute value of the estimated residual.

I use two different types of tests for testing long-run relationships. The first one, which I use more intensively, is Pesaran (1997) and Pesaran and Shin's (1999) ARDL approach to long-run modeling. This test has the nice feature that it estimates consistent long-run relationship independently of whether the variable of interest are integrated or stationary. The second one is Johansen (1988) and Johansen and Juselius' (1990) cointegration test, which estimates long-run relationships among only integrated series. I use the latter as supporting evidence in the first set of estimates. After this, I stick to the Pesaran and Shin's (1999) methodology because of the known problems in determining the nonstationary characteristics of time series. [5] [6]

[5] See Loayza and Ranciere (2006) for an excellent description of the ARDL approach to long-run modeling.

[6] The ARDL results are based on a ARDL(2,1) which implies the following error correction model; y_t represents the level of aggregate volatility in period t and x_t represents the level of idiosyncratic volatility.

$$\Delta y_t = \gamma_1 \Delta y_{t-1} + \gamma_2 \Delta x_t + \alpha[y_{t-1} - \beta_0 - \beta_1 x_{t-1}] + \mu_t$$

The cointegration analysis is based on an error correction model with one lag; $z_t = [y_t \ x_t]$.

$$\Delta z_t = \delta_1 \Delta z_{t-1} + \alpha\big[y_{t-1} - \beta_0 - \beta_1 x_{t-1}\big] + \nu_t$$

I use the test proposed in Bai and Perron (1998, 2003) to test for structural breaks in each series; this test estimate an *optimal* break-date. I complement the analysis with hypothesis tests on the null hypothesis of joint-contemporaneous breaks in both series. The details of the hypothesis tests are described in Appendix 2.A.3. [7]

2.4.1 Full sample

2.4.1.1 Long-run relationship

Table 1 presents the results of the tests on the long-run relationship between aggregate volatility and idiosyncratic volatility. The period covered by the estimation is 1951-2004. The main conclusions are summarized in Empirical Result 1.

Empirical Result 1. *I reject the null hypothesis of no long-run relationship between aggregate and idiosyncratic volatility (full sample estimate). The estimated negative long-run relationship goes in the direction predicted by both types of theories, competition-based and diversification-based theories. The magnitude of the long-run coefficients implies that a 100 basis point increase in idiosyncratic volatility is associated with approximately a 11 basis point lower aggregate volatility.* [8]

[7] I am going to assume that the series present at most one structural change.

[8] A likelihood ratio test is used to test for the significance of the long-run parameters in the cointegration equation (part 2 in Table 1). This test is performed by comparing the likelihood of the error correction model estimated for Δy_t (unrestricted model) and the likelihood of the model $\Delta y_t = \lambda_1 + \lambda_1 \Delta y_{t-1} + \lambda_2 y_{t-1} + v_t$ estimated by OLS (restricted model).

Table 1: Long-Run Relationship between Idiosyncratic and Aggregate Volatility (1951-2004)

	Idiosyncratic Volatility		Aggregate Volatility
1.- ARDL approach to long-run modeling (Pesaran–Shin–Smith):			
Long-run equation coefficients	-0.110		1.00
H_0: $\beta = 0$ (SE)	(0.022)**		
2.- Cointegration analysis (Johansen–Juselius):			
H_0: None CE (LR-statistic)		39.29**	
H_0: At most 1 CE (LR-statistic)		3.56	
Cointegration coefficients	-0.115		1.00
H_0: $\beta = 0$ (LR-Test dof=2)	14.25**		

2.4.2 Structural break

Table 2 presents the result for the structural break tests; the break-date is defined as the last period of the old regime.

Empirical Result 2. *I reject the null hypothesis of no structural change for both series. The estimated optimal break is 1984 for the aggregate volatility series and 1980 for the idiosyncratic volatility series (full sample estimate). I reject the null that the break in aggregate volatility occurred in 1980 against the alternative that it occurred in 1984. I also reject the null that the break in idiosyncratic volatility occurred in 1984 against the alternative that it occurred in 1980. As well, I reject the null that both series had a joint-contemporaneous break in 1980, or in 1984, against the alternative that the breaks occurred at the optimal dates (idiosyncratic volatility break in 1980 and aggregate volatility break in 1984).*

Table 2: Structural Break (1951-2004)

	Idiosyncratic Volatility		Aggregate Volatility
1.- Univariate structural break analysis I (Bia–Perron):			
Test of structural break (*sup*F-statistic)	33.59**		20.19**
Estimated year of break	1980		1984
2.- Univariate structural break analysis II:			
H_0: 1980 Break ; H_1: 1984 Break ‡	—		4.36*
H_0: 1984 Break ; H_1: 1980 Break ‡	15.41**		—
‡(LR-Test dof=1)			
3.- Test of contemporaneous break (d):			
Year of Break under H_0	1984		1984
Year of Break under H_1	1980		1984
Log-Likelihood Ratio Test (dof=1)		10.02**	
Year of Break under H_0	1980		1980
Year of Break under H_1	1980		1984
Log-Likelihood Ratio Test (dof=1)		6.22*	

Note: (a) Statistic in parenthesis; (b)* rejects H_0 at 5%, ** rejects H_0 at 1%; (c) LR: likelihood ratio; (d) aggregate and idiosyncratic volatility are standardized using the formula $\frac{X_{i,t}-\bar{X}_i}{S_i}$ for i =aggregate, idiosyncratic; (e) critical values for $\chi^2_{(1,0.05)} = 3.84$ and $\chi^2_{(1,0.01)} = 6.63$.

2.4.3 Sample issues

The empirical literature in various fields has identified a potential sample problem of time series aggregates based on the Compustat sample. The continuing entry of new firms into the stock market implies that the estimate of $\hat{g}^{\mathrm{ID}}_{i,t}$ could be driven by changes in the sample composition.

I study this issue with a simple, but enlightening exercise. In order to construct a more stable sample from where to draw conclusions about the evolution of firms' idiosyncratic volatility, I estimate $\hat{g}^{\mathrm{ID}}_{i,t}$ for two disjoint subsamples: i) each year's top-500 firms according to their level of sales and ii) the rest of the firms in the stock market. Panel A in Figure 2.4 shows the evolution of the idiosyncratic volatility for each subsample and for the whole sample as well. Panel B shows the series for the top-500 firms and a filtered version (± 3 years moving average).

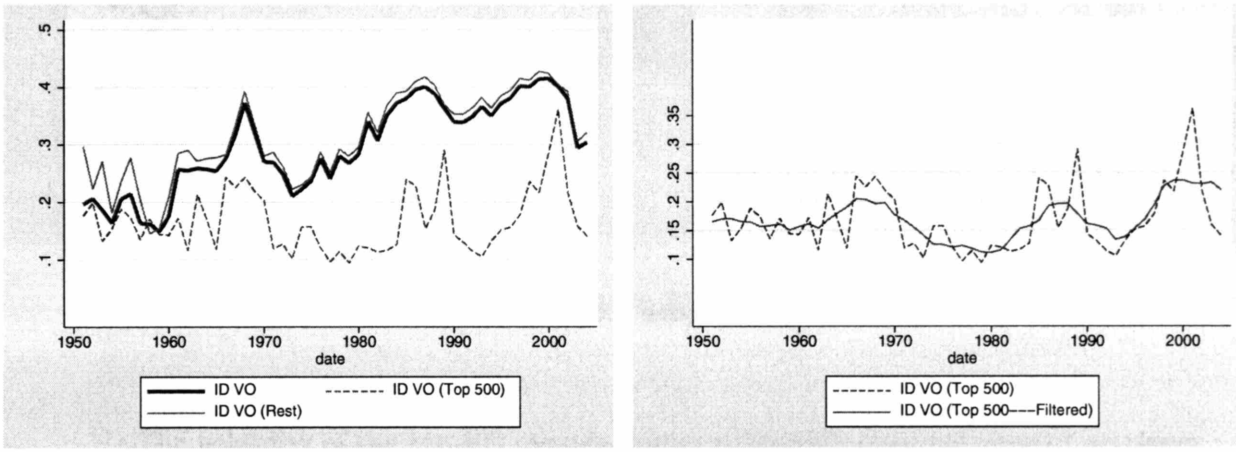


Figure 2.4.

I apply Pesaran and Shin's long-run relationship test to study the relationship between aggregate volatility and idiosyncratic volatility of the top-500 firms. The coefficient of the top-500 idiosyncratic volatility variable in the long-run equation is -0.050 with a standard error of 0.057 (insignificantly different from zero).

Empirical Result 3. *The increased idiosyncratic volatility suggested by the analysis of the full sample, does not apply uniformly across different types of firms. In particular, the idiosyncratic volatility has not increase among the top-500 firms during the last 5 decades. I cannot reject the null hypothesis that there is no long-run relationship between aggregate volatility and idiosyncratic volatility of the top-500 firms.*

Figure 2.5 studies the stability of the subsample containing the top-500 firms. Panel A plots the share of total sales in the full sample represented by each subsample. Panel B analyzes the survival rate of firms among the top-500 subsample. The top line is the number of firms in the top-500 firms subsample that are also present in the top-500 subsample one year before; the middle (lower) line does a similar analysis but for firms present in the top-500 subsample 5 (10) years before. In 10 years, more than half of the firms still remain in the top-500 subsample.

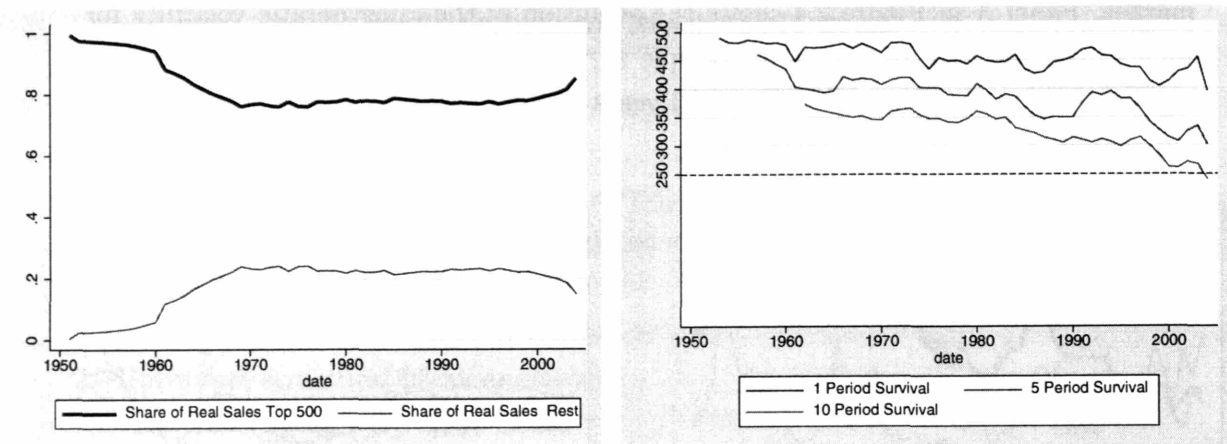
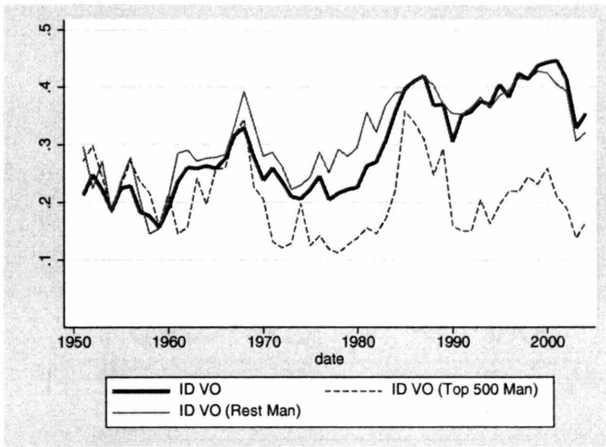


Figure 2.5.

Finally, I compute idiosyncratic volatility based on the top-500 subsample within manufacturing firms. This sector is more homogeneous than the full sample, and although its share of total sales has decreased over time, it continues to represent at least 35% of all US listed firms. Similarly, it is expected that this subsample contains fewer holding-type firms. This type of firms can distort the top-500 analysis because they tend to concentrate a more diversified portfolio of production process than producing-type firms. The results for the manufacturing subsample are shown in Figure 2.6 and they also suggest that the volatility of all individual firms has not increased during the last decade. The survival characteristics of the top-500 subsample for the manufacturing sector are very similar to the ones pictured in panel B of Figure 2.5.



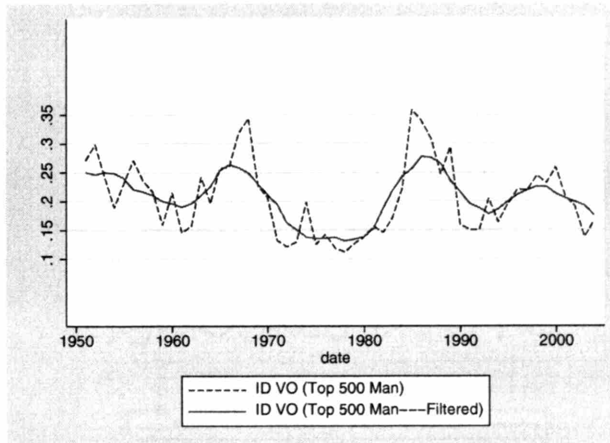


Figure 2.6.

The volatility of the top-500 manufacturing subsample does not present an upper trend and is consistent with the results for the top-500 firms.

2.4.4 Idiosyncratic volatility of new firms

The subsample of non-top-500 firms is characterized by much younger firms than the top-500 subsample—I define year 0 as the year in which the firm entered the market. It is well known that new firms tend to be more volatile than more mature firms.

Figure 2.7 shows that the number and importance of new firms has not been increasing throughout the last decades. Nor there has been any sharp changes in the number and importance of new firms around the year 1984 (year of the Great Moderation; vertical line).

I apply Pesaran and Shin's long-run relationship test to study the relationship between idiosyncratic volatility (full sample) and number of new firms entering the stock market. The coefficient of the number of new firms in the long-run equation is 3.09 with a standard error of 5.82 (insignificantly different from zero). If we use the importance of new firms (sales) instead of the number of new firms, the coefficient in the long-run equation is -3.96 with a standard error of 4.83 (insignificantly different from zero).

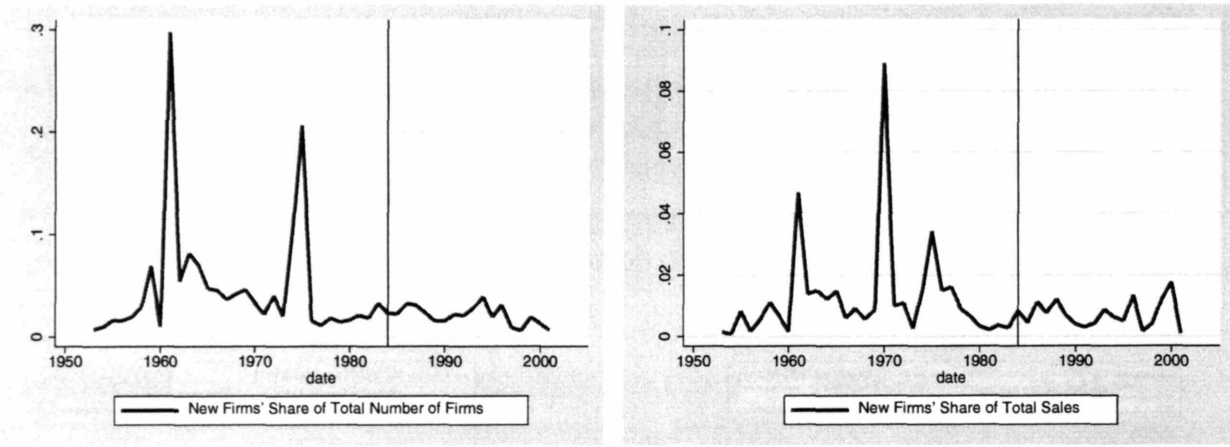


Figure 2.7. New firms in period t are defined as firms for which the first available observation in Compustat is in period t. The share is over the total number of firms in period t and sum of all firms' sales in period t, respectively.

Although the number and importance of new firms did not change significantly during the last 5 decades, the characteristics of the new firms entering the market did change during the last decades. This evolution is characterized in Figures 2.8 and 2.10. Table 3 presents formal analysis regarding the long-run relationship between new firms' characteristics and idiosyncratic volatility (full sample).

Figure 2.8 shows time series for the idiosyncratic volatility of subsamples of firms according to the period in which these fims entered the stock market. For example, the thick line in the left panel represents the idiosyncratic volatility of the subsample of firms that entered the market between 1962 and 1969. Period 0 represents the last year of the cohort (*e.g.*: 1969 for the cohort 1962-1969, 1985 for the cohort 1978-85, etc.). Figure 2.9 shows the relative importance of each cohort across time as measured by their share of total sales.

The idiosyncratic volatility of each cohort tends to decrease with time, as expected. Across cohort, the picture reveals an increase in the idiosyncratic volatility of new firms entering after 1977. Figure 2.10 studies this in more detail.

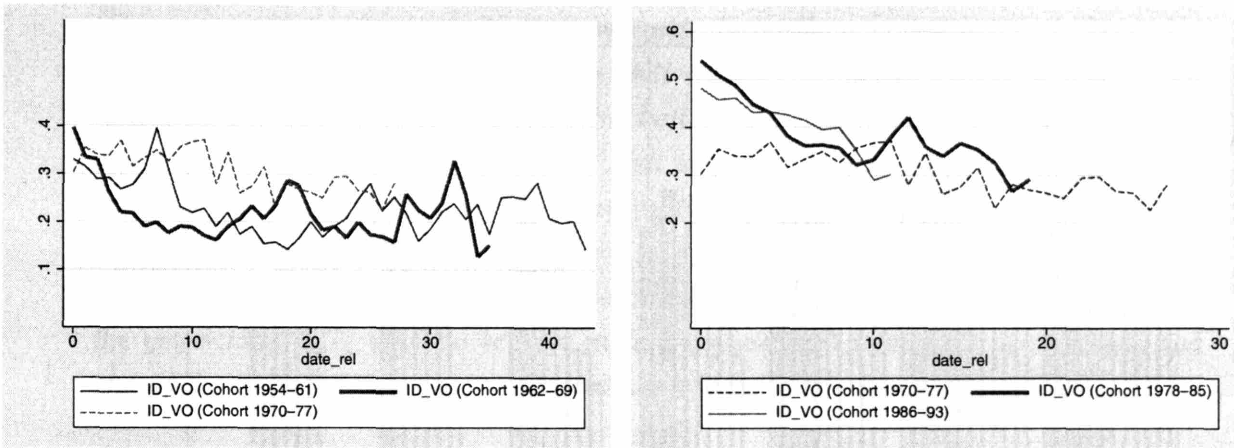


Figure 2.8.

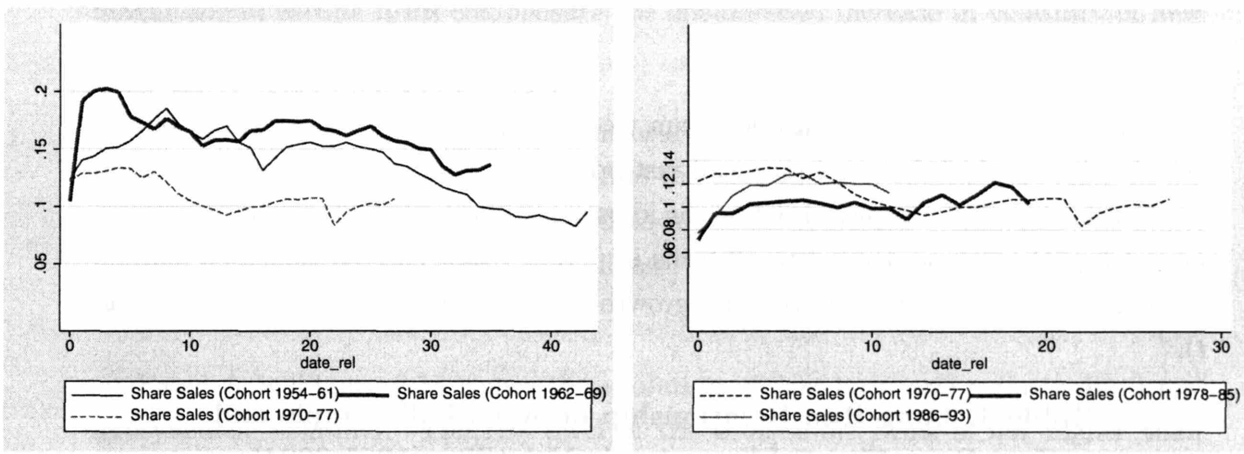


Figure 2.9.

The sample cohorts in Figure 2.10 have a smaller range of 4 years instead of 8 years. Each bar is the average idiosyncratic volatility observed between the years indicated in the X-axis for the corresponding cohort. The shaded columns correspond to cohorts of firms entering the stock market after 1977. During the first 4 years in the market, the volatility of firms entering after 1977 is 200 basis points higher than the volatility in the first 4 years of firms entering before 1978 (+60% higher). Even after being 15 years in the stock market, the difference in volatility is no less than 130 basis points (+60% higher).

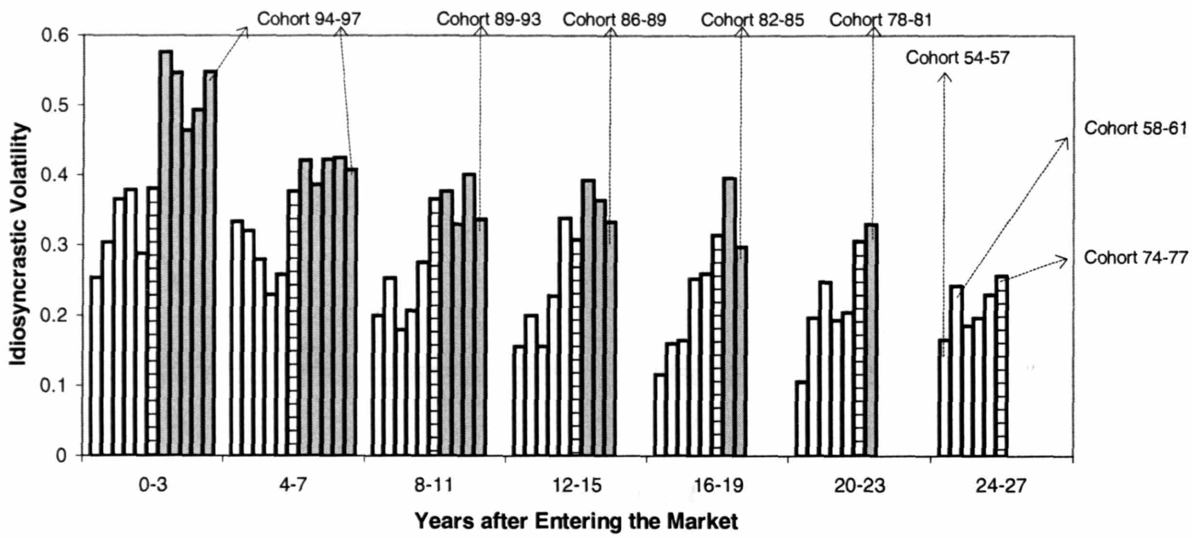


Figure 2.10.

Table 3 formally tests for the long-run relationship between the volatility of the entering firms, which I am going to call just-born firms, and the measures of idiosyncratic volatility (full sample). I define the idiosyncratic volatility of just-born firms in period t as the risk associated to all firms in period t that entered the stock market on period $t-3$, $t-2$ or $t-1$ (the first growth observation is in period $t-2$, $t-1$ or t).

Table 3: Long-Run Relationship between Full Sample and Just-Born Firms' Idiosyncratic Volatility (1955-2004)

	Idiosyncratic Volatility	Just-Born Volatility
1.- ARDL approach to long-run modeling (Pesaran–Shin–Smith):		
Long-run equation coefficients	1.00	0.530
H_0: $\beta = 0$ (SE)		(0.104)**
2.- Univariate structural break analysis (Bia–Perron):		
Test of structural break (*sup*F-statistic)	33.59**	29.56**
Estimated year of break	1980	1979

Note: (a)* rejects H_0 at 5%, ** rejects H_0 at 1%; (b) SE: Standard error.

Empirical Result 4. *During the last 5 decades, there is no long-run relationship* *between the number, or importance (in sales), of new firms entering the stock market and the level of idiosyncratic volatility (full sample). However, there is a long-run relationship between the level of risk implicit in the new firms entering the market and the level of idiosyncratic volatility (full sample). In fact, the firms that entered the stock market after 1977 show a significantly higher level of idiosyncratic volatility compared to the firms that entered before 1974. This difference lasts at least 15 years after the firms entered the market. The cohort of firms entering the market between the years 1974-1977 seem to be in a middle ground with low levels of risk until the end of the 70's and then catching-up with the firms that entered after 1977.*

We have concentrated so far in the evolution of the new firms. However, the picture would not be complete if we did not have a measure of the implicit risk associated to the firms that are exiting the market. It could be that the risk and the magnitude of exiting firms compensate the documented increase in volatility of new firms.

To address this issue, I estimate a measure of the implicit associated risk of just-born firms and soon-to-die firms. I define the idiosyncratic volatility of soon-to-die firms in period t as the risk associated to all firms in period t that exit the stock market on period t, $t+1$ or $t+2$. I use the same definition of just-born firms as before. [9]

Panel A in Figure 2.11 shows the evolution of just-born and soon-to-die firms' idiosyncratic volatility. The risk level of the soon-to-die firms is not higher than the risk associated to the just-born firms. In fact, after the mid 70's a positive gap appears in favor of the just-born firms. Also, I compute the importance of the total sales in each subgroup with respect to the total sales of all the firms in the sample. Panel B in Figure 2.11 shows that the amount of sales represented by the soon-to-die firms has never been significantly higher than the sales of the just-born firms, except around the year 2000. [10]

[10]The total sales series for the soon-to-die firms in Panel B of Figure 2.11 has been lagged 3 periods. This is done to create a more accurate picture of the "magnitude of sales" entering and exiting the market in each period.

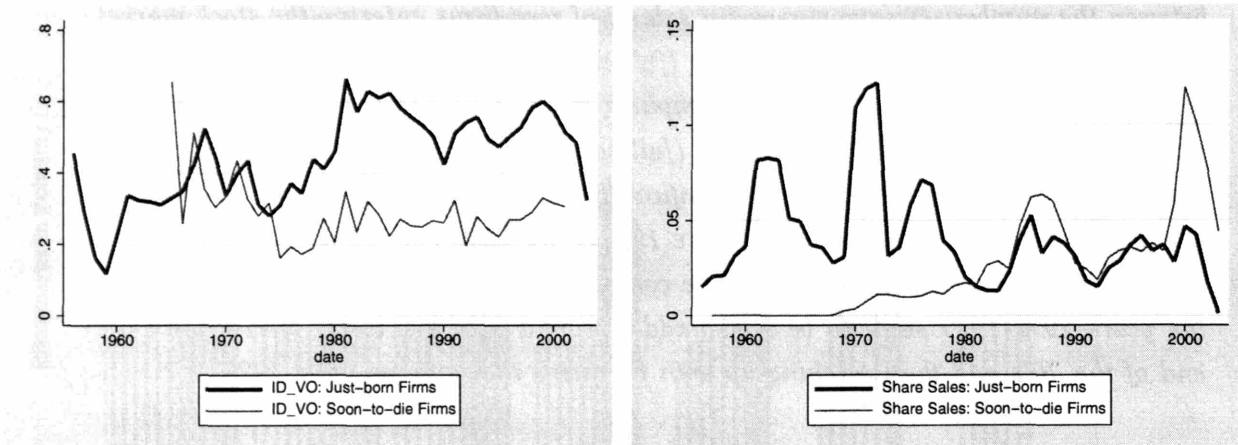


Figure 2.11.

Empirical Result 5. *The idiosyncratic volatility and the importance (in sales) of exiting firms does not compensate the increase in the volatility of new firms entering the stock market.*

2.4.5 Macro stabilization policies

Finally, I study the long run relationship between aggregate and idiosyncratic volatility controlling for changes in the macro policy variables. Changes in the inflation rate and in fiscal behavior can be potential explanations of the great moderation since they can affect the macro stability of the economy.

Table 4 show the estimates for the long-run equation. The fiscal volatility variable is computed in a similar fashion to aggregate volatility using real government expenditures growth rate deflated by the CPI. Inflation rate is computed using the CPI.

Table 4: Long-Run Relationship between Idiosyncratic, Aggregate Volatility, and policy variables (1951-2004)

	Idiosyncratic Volatility	Inflation	Fiscal Volatility	Aggregate Volatility
1.- ARDL approach to long-run modeling (Pesaran–Shin–Smith):				
Long-run equation coefficients	-0.113	0.096	0.004	1.00
H_0: $\beta = 0$ (SE)	(0.021)**	(0.0561)*	(0.161)	

Note: (a)* rejects H_0 at 5%, ** rejects H_0 at 1%; (b) SE: Standard error.

Empirical Result 6. *When controlling for the level of inflation and the level of fiscal volatility, the estimated long run relationship between aggregate and idiosyncratic volatility still holds. The coefficient of the inflation variable is statistically significant (at 5%), while the coefficient of fiscal volatility is insignificantly different from zero.*

2.5 Conclusions

This paper contributes to the empirical literature on the relationship between firm-level micro volatility and macro aggregate volatility.

The main results reveal that there has been an increase in the average volatility of US listed firms during the last decades, and that it is negatively related in the long run to aggregate volatility. The increase in idiosyncratic volatility has not occurred in all types of firms. In particular, the top-500 firms (according to sales) do not show any trend nor evidence of a long-run relationship with aggregate volatility during the last 5 decades.

Regarding new firms, there is no trend in the number and importance of new firms entering the stock market—we cannot associate the increase in idiosyncratic volatility with more new firms in the market. What we do observe, however, is a significantly higher idiosyncratic volatility of new firms entering the stock market at the end of the 70's—not more new firms, but more risky new firms. This increase is not compensated by an increase in risk and magnitude of firms exiting the stock market.

Finally, when controlling for macro policies—inflation and fiscal volatility—the long run relationship between aggregate and idiosyncratic volatility still holds.

2.A Appendix

2.A.1 Consistency of $\hat{\sigma}_t^{2\ \mathrm{ID}}$

Let's analyze separately each of the 3 terms in equation (2.12).

First term:

With respect to the first term, we know from (2.5) that $\mathbb{E}\,[g_{i,t}^{\mathrm{ID}}]^2 = \sigma_t^{2\ \mathrm{ID}}$. Applying the law of large numbers, we can state that

$$\frac{1}{N_t}\sum_{\forall i,s}[g_{i,s,t}^{\mathrm{ID}}]^2 \xrightarrow{p} \sigma_t^{2\ \mathrm{ID}} \tag{2.14}$$

Second term:

With respect to the second term, we can rewrite

$$\left(g_{i,s,t}^{\mathrm{NID}} - \hat{g}_{i,s,t}^{\mathrm{NID}}\right) = \left(g_t + g_{s,t} + g_{i,t}^{\mathrm{IE}}\right) - \left(\hat{g}_t + \hat{g}_{s,t} + \hat{g}_{i,t}^{\mathrm{IE}}\right) \tag{2.15}$$

and

$$\hat{g}_t + \hat{g}_{s,t} = \frac{1}{N_{s,t}}\sum_{i=1}^{N_{s,t}} g_{i,s,t} = \frac{1}{N_{s,t}}\sum_{i=1}^{N_{s,t}}\left(g_t + g_{s,t} + g_{i,t}^{\mathrm{IE}} + g_{i,t}^{\mathrm{ID}}\right) =$$

$$(g_t + g_{s,t})\frac{1}{N_{s,t}}\sum_{i=1}^{N_{s,t}} + \frac{1}{N_{s,t}}\sum_{i=1}^{N_{s,t}}(g_{i,t}^{\mathrm{IE}} + g_{i,t}^{\mathrm{ID}}) = \tag{2.16}$$

$$(g_t + g_{s,t}) + \frac{1}{N_{s,t}}\sum_{i=1}^{N_{s,t}}(g_{i,t}^{\mathrm{IE}} + g_{i,t}^{\mathrm{ID}})$$

From equation (2.2) we know that $\frac{1}{N_{s,t}}\sum_{i=1}^{N_{s,t}}(g_{i,t}^{\mathrm{IE}}+g_{i,t}^{\mathrm{ID}}) \xrightarrow{p} 0$ and $\frac{1}{n_{s,t}}\sum_{i=1}^{n_{s,t}}(g_{i,t}^{\mathrm{IE}}+g_{i,t}^{\mathrm{ID}}) = 0$. Therefore,

$$\hat{g}_t + \hat{g}_{s,t} \approx g_t + g_{s,t} \tag{2.17}$$

As well, we can write

$$\hat{g}_i^{\mathrm{IE}} = \frac{1}{2w+1} \sum_{r=t-w}^{t+w} \left((g_r + g_{s,r} + g_{i,r}^{\mathrm{IE}} + g_{i,r}^{\mathrm{ID}}) - \hat{g}_r - \hat{g}_{s,r} \right) =$$

$$\frac{1}{2w+1} \sum_{r=t-w}^{t+w} g_{i,r}^{\mathrm{IE}} + g_{i,r}^{\mathrm{ID}} + \frac{1}{2w+1} \sum_{r=t-w}^{t+w} \left(g_r + g_{s,r} - \hat{g}_r - \hat{g}_{s,r} \right) = \tag{2.18}$$

$$g_{i,t}^{\mathrm{IE}} + \frac{1}{2w+1} \sum_{r=t-w}^{t+w} \left(g_r + g_{s,r} - \hat{g}_r - \hat{g}_{s,r} \right)$$

Based on the latter formulation, when $N_t \to \infty$ we can write

$$\left(g_{i,s,t}^{\mathrm{NID}} - \hat{g}_{i,s,t}^{\mathrm{NID}} \right) \approx 0 \tag{2.19}$$

and

$$g_{i,t}^{\mathrm{ID}} \cdot \left(g_{i,s,t}^{\mathrm{NID}} - \hat{g}_{i,s,t}^{\mathrm{NID}} \right) \approx 0 \tag{2.20}$$

and

$$\mathbb{E} \left[g_{i,t}^{\mathrm{ID}} \cdot \left(g_{i,s,t}^{\mathrm{NID}} - \hat{g}_{i,s,t}^{\mathrm{NID}} \right) \right] \approx 0 \tag{2.21}$$

Applying the law of large numbers, we can state that

$$2 \frac{1}{N_t} \sum_{\forall i,s} [g_{i,s,t}^{\mathrm{ID}}] \left(g_{i,s,t}^{\mathrm{NID}} - \hat{g}_{i,s,t}^{\mathrm{NID}} \right) \xrightarrow{p} 0 \tag{2.22}$$

Third term:

With respect to the third term, remember that $\left(g_{i,s,t}^{\mathrm{NID}} - \hat{g}_{i,s,t}^{\mathrm{NID}} \right) \approx 0$ and $\mathbb{E} \left(g_{i,s,t}^{\mathrm{NID}} - \hat{g}_{i,s,t}^{\mathrm{NID}} \right) \approx 0$. As a result,

$$\frac{1}{N_t} \sum_{\forall i,s} \left(g_{i,s,t}^{\mathrm{NID}} - \hat{g}_{i,s,t}^{\mathrm{NID}} \right)^2 \xrightarrow{p} 0 \tag{2.23}$$

Finally, based on equations (2.14), (2.22), and (2.23) we can state that

$$\lim_{N_t \to \infty} \hat{\sigma}_t^{2\ \mathrm{ID}} = \sigma_t^{\mathrm{ID}} \qquad \left[\hat{\sigma}_t^{2\ \mathrm{ID}} \xrightarrow{p} \sigma_t^{\mathrm{ID}}\right] \tag{2.24}$$

2.A.2 Sample composition

Tale A1: Sample Composition

Year	Number of Firms	Year	Number of Firms
1951	576	1978	5,278
1952	714	1979	5,104
1953	721	1980	5,038
1954	728	1981	4,990
1955	747	1982	4,979
1956	759	1983	5,036
1957	777	1984	5,044
1958	800	1985	5,038
1959	866	1986	5,102
1960	908	1987	5,329
1961	1,687	1988	5,399
1962	1,841	1989	5,392
1963	2,088	1990	5,427
1964	2,344	1991	5,475
1965	2,515	1992	5,511
1966	2,701	1993	5,594
1967	2,876	1994	6,061
1968	3,069	1995	6,163
1969	3,299	1996	6,391
1970	3,410	1997	6,450
1971	3,487	1998	6,106
1972	3,633	1999	6,152
1973	3,729	2000	6,121
1974	3,916	2001	6,042
1975	5,255	2002	5,628
1976	5,339	2003	5,042
1977	5,267	2004	3,893
		Total	211,837

Note: Number of firms in each period in the sample used to compute the results presented in this paper.

2.A.3 Hypothesis tests of structural break

The first set of estimates in Table 2 tests the null hypothesis of no structural break against the hypothesis of a structural beak at an unknown period. These tests are based on Bia and Perron (1998, 2003) and they also estimate the optimal break-period.

The second set of hypotheses tests studies the null hypothesis (H_0) that each univariate series had a structural break in the period where the other series had its optimal break. The null hypothesis is tested against the alternative hypothesis (H_1) that the break occurred in the optimal period estimated for each series. These hypotheses are tested with log likelihood ratio tests based on the following model.

$$y_t = \theta_1 + \lambda_1 D_{84} + \lambda_2 D_{80} \tag{2.25}$$

where y_t is the corresponding series (either aggregate or idiosyncratic volatility); D_T is a dummy variable that takes the value of 1 if $t > T$. For the idiosyncratic volatility series, the null hypothesis corresponds to the case where $\lambda_2 = 0$, while the alternative corresponds to the case where $\lambda_1 = 0$. The opposite configuration holds for the aggregate volatility series.

The third set of estimates tests the null hypothesis of joint-contemporaneous break—that both series had a common break either in 1980 or 1984—against the alternative that each series had its break in their estimated optimal break date. The null hypothesis that both series had a break in 1984 is tested with a log likelihood ratio tests based on the following model:

$$z_{it} = \theta_1 D_A + \theta_2 D_I + \lambda_1 D_A D_{84} + \lambda_2 D_I D_{84} + \lambda_3 D_I D_{80} + \mu \tag{2.26}$$

where z_{it} is a variable with the aggregate (A) and idiosyncratic (I) variables stacked (for $i = A, B$); D_i is a dummy variable that takes a value of 1 for the observations corresponding to the series i (for $i = A, B$); D_T is a dummy variable that takes the value of 1 in periods $t > T$. The $\lambda_3 = 0$ case corresponds to the null hypothesis that both breaks occurred in 1984, while the $\lambda_2 = 0$ case corresponds to the alternative hypothesis that the breaks occurred at each individual optimal date.

The test for the null hypothesis of joint structural break in 1980 is analogous and it is based on the following model:

$$z_{it} = \theta_1 D_A + \theta_2 D_I + \lambda_1 D_A D_{80} + \lambda_2 D_I D_{80} + \lambda_3 D_A D_{84} + \mu \tag{2.27}$$

where $H_0: \ \lambda_3 = 0$ and $H_1: \ \lambda_1 = 0$.

Finally, the variables in z_{it} are standardized using the formula $\frac{x_{i,t} - \bar{x}_i}{S_i}$ for $i = A, B$, where $S = \frac{1}{N}\sum_{j=1}^{N}(x_{i,t} - \bar{x}_i)^{0.5}$. If we do not standardized the measures of volatility, then difference in the standard deviation of the series will mechanically bias the result in favor to the optimal date of the variable with higher variance. In this case, in favor of the optimal date of the idiosyncratic volatility series that has a standard deviation almost 4 times higher than the standard deviation of the aggregate volatility series. In fact, when I do the test under no standardization, I cannot reject the null of a joint-contemporaneous break in 1980 (p-value of 28%).

Chapter 3

Fiscal and Monetary Policy Biases under Lack of Coordination

> "The most recent [Polish] parliament added recklessly to spending. This has led to an awkward stand-off with the central bank, whose governor, Leszek Balcerowicz, was the father of shock therapy when he was finance minister in 1990. Mr Balcerowicz and his colleagues at the bank say interest rates will have to stay high ... until the politicians can demonstrate fiscally responsible behaviour. Ministers complain that the central bank is strangling growth." (*Limping towards Normality*, Oct. 25th, 2001, The Economist.)

Until not so long ago, the debate on the relationship between monetary and fiscal authorities has been centered on the inflationary consequences of the monetary financing of the fiscal deficit. The moderately high inflation of the 1970s in some industrialized countries and, particularly, the recurring episodes of very high inflation in several developing countries justified this focus. The main policy recommendation to avoid high and variable inflation has been the institution of an independent monetary authority whose main mandate would be to ensure price stability (see Cukierman 1992, and Walsh 1993). In fact, in recent years several central banks have adopted inflation targeting as the cornerstone of their monetary policy (see Bernanke *et al.* 1999).

A second policy recommendation has been to impose discipline in the fiscal accounts and reduce the public deficit. This has been achieved by both rationalizing fiscal expenditures (e.g., eliminating price subsidies and privatizing public enterprises) and raising tax revenues, particularly through the adoption of value-added taxes. Furthermore, fiscal authorities are using domestic and international financial markets to better manage the public debt to avoid the need to collect seignorage from outstanding monetary assets.

Thus, in many countries around the world there is a new policy environment, one in which monetary authorities are committed to controlling inflation and fiscal authorities do not rely on the inflationary tax to finance their deficits and debt service. In this new context, a different set of policy issues and questions arise. This paper is devoted to the study of one of the most important of them, namely, the effect of the lack of coordination between monetary and fiscal authorities in achieving the goals of minimizing business cycle fluctuations.

Coordination (or the lack thereof) is a relevant issue because the monetary and fiscal authorities have different policy instruments, different objectives and preferences, and sometimes different perceptions of how the economy functions. These differences can lead both authorities to behave antagonistically and produce undesirable results in the economy. Following Nordhaus (1994) and Loewy (1988), we use a game-theoretic approach to analyze the effects of the lack of policy coordination between fiscal and monetary authorites on public deficits and domestic real interest rates. We then test the main implications of the model using a sample of 19 industrial countries over the period 1970-94.

In order to illustrate the main ideas of the paper, let's consider a simple fiscal-monetary game modeled after the well-known "prisoner's dilemma." Figure 1 presents the main assumptions and results of this game. We analyze the potential response of the monetary and fiscal authorities in the face of a negative shock that rises inflation and lowers employment. The monetary and fiscal authorities have two options to confront a negative supply shock: they can either follow loose or tight policies. When both "play" tight, the resulting inflation is low but so is the resulting employment. When both play loose, both inflation and employment are high. And when only one of them plays tight, the result is medium employment and inflation.

The distinctive feature of this fiscal/monetary game is that monetary and fiscal authorities have different preferences for inflation and employment (see the payoff schedules in Figure 1). Whereas the monetary authority considers more valuable to achieve low inflation than high employment, the fiscal authority regards obtaining high employment as more important than keeping inflation low. The preference dif-

Figure 1: A Monetary-Fiscal Game

		Central Bank	
		Tight Monetary	Loose Monetary
Fiscal Authority	Tight Fiscal	$\boxed{7}$ $\boxed{4}$ Low Inflation Low Employment	$\boxed{6}$ $\boxed{6}$ Medium Inflation Medium Employment
	Loose Fiscal	$\boxed{6}$ $\boxed{6}$ Medium Inflation Medium Employment	$\boxed{4}$ $\boxed{7}$ High Inflation High Employment

Payoff Schedules

Inflation	Low	Medium	High
Central Bank	6	4	1
Fiscal Authority	3	2	1

Employment	Low	Medium	High
Central Bank	1	2	3
Fiscal Authority	1	4	6

ferences between both authorities are chosen to be sufficiently large so as to obtain the result we would like to stress.

The only Nash equilibrium in this game consists of a tight monetary policy and a loose fiscal policy. The other three alternatives present opportunities for one of the players to benefit by unilaterally deviating from the original play. The equilibrium of this game exposes the paradigmatic conservatism of central banks and liberalism of fiscal authorities. It also illustrates why their responses are optimal given the preference differences between the two. For instance, if the monetary authority were to follow a loose policy –believing fiscal authority's assurance for stricter restraint–, then the fiscal authority would find it optimal to renege its pledge and play a loose policy. Note that in terms of the payoffs to both authorities, the Nash equilibrium is equivalent to the combination of loose monetary and tight fiscal policies. From a long-run perspective, it can be argued that the latter combination of policies is healthier than the Nash equilibrium given that it does not compromise fiscal sustainability and does not weaken the investment capacity of the private sector.

In the second section of the paper we formalize these ideas through a monetary-fiscal game, clarifying the conditions under which looser fiscal policy (represented by higher primary fiscal deficits) is accompanied by tighter monetary policy(represented by higher real interest rates), as predicted by the "prisoner's dilemma" game. The basic conclusion of the model is that if the preferences for output and inflation gaps become more dissimilar between the monetary and fiscal authorities, then the primary fiscal deficit and real interest rate will rise by a larger amount in the face of economic shocks.

Also in the theoretical section, we compare the Nash equilibrium solution with the Stackelberg solution. By allowing one of the authorities to be the leader, the Stackelberg solution introduces dynamic aspects into the game. It creates the possibility for the leading authority to act in a way that elicits a mutually beneficial response from the follower. The Stackelberg game gives the same qualitative conclusion as the Nash equilibrium; however, due to its potential for beneficial interaction, the Stackelberg equilibrium renders lower policy biases –that is, lower fiscal deficits and interest rates in the face of shocks.

The third section of the paper bring some empirical evidence to bear. We use annual information over the period 1970-94 for a sample of industrial countries to test the main conclusion of the paper: in a context where fiscal and monetary authorities are independent and do not effectively coordinate their policy responses, countries where the fiscal and monetary authorities are more divergent regarding their preferences for output and inflation gaps will exhibit larger primary deficits and real interest rates. Given the highly simplified nature of our game-theoretic model, this conclusion would apply only after controlling for other factors affecting the level of primary deficits and domestic real interest rates.

We run a system of reduced-form regressions with the primary deficit (as ratio to GDP) and the real domestic interest rate (as deviation from the international interest rate) as dependent variables. We proxy the central bank's relative preference in favor of reducing inflation deviations by an indicator of the importance of price stability in the central bank's charter. Likewise, we proxy the fiscal authority's relative preference in favor of reducing the output gap by an index of political orientation of the party in power. In these regressions we control for a number of factors that may be correlated with the dependent variables, such as business cycle effects, international conditions, and Ricardian-equivalence effects. We find evidence robustly consistent with the main conclusion of the theoretical model.

The policy implication we derive from these results is that, notwithstanding the gains from central bank independence, there are gains to be made from policy coordination between monetary and fiscal authorities. However, this implication would apply only to countries that have already achieved price stability and fiscal discipline.

3.1 A Game-Theoretic Model

This section presents a one-period game played by monetary and fiscal authorities. It builds from the trade-off that the economy faces in the short-run between changes in the inflation rate and the output gap (Phillips Curve). The model emphasizes the effects on the level of fiscal deficits and real interest rates that result from different preferences by the monetary and fiscal authorities with respect to inflation and output deviations from their optimal level.

This game-theoretic approach is based on Frankel (1988), Loewy (1988) and Nordhaus (1994). The main difference between Frankel's model and ours is that Frankel assumes a world where the authorities have the same preference with respect to inflation and output deviations but disagree on the model that best represents the economy. [1] Our model is similar to Nordhaus', with the difference that we assume asymmetric preferences for negative and positive changes of inflation and output, we allow for both authorities to be reluctant to change their policy instruments, and we compare the Nash and Stackelberg equilibria results.

[1] Frankel concludes that policy coordination may not be welfare improving if it means a departure from the "true" model. However, coordination would be more likely to be welfare improving if it means sharing information and agreeing on a common model for the economy.

3.1.1 The model

We assume that policy makers seek to maximize an asymmetric utility function. Both fiscal and monetary authorities dislike a fall in output and a rise in inflation; however, they do not mind the opposite. In addition, we assume that both authorities dislike changing their respective policy instrument from its long-run level. [2]

The fiscal authority's utility function is denoted by U^F and the unnormalized preference weights for each objective are given by the coefficients α_u^F, β_u^F, and δ_u (equation 3.1). They measure respectively the cost associated to output (y) falling below a certain threshold (y^*), to inflation (π) rising beyond its long-run level (π^*) and to deficit deviations from its long-run level ($D - D^*$). Note that α_u^F, β_u^F, $\delta_u > 0$. Given the ordinal nature of the utility function and in order to clarify the role of the assumptions, let's normalize the preference parameters so that they are given with respect to the cost of inflation deviations: $\alpha^F = \alpha_u^F/\beta_u^F$ and $\delta = \delta_u/\beta_u^F$ (equation 3.2).

$$U^F = U^F\{(y - y^*), (\pi - \pi^*), (D - D^*)\}$$

$$U^F = -\alpha_u^F\{\min(y - y^*, 0)\}^2 - \beta_u^F\{\max(\pi - \pi^*, 0)\}^2 - \delta_u(D - D^*)^2 \tag{3.1}$$

$$U^F = -\alpha^F\{\min(y - y^*, 0)\}^2 - \{\max(\pi - \pi^*, 0)\}^2 - \delta(D - D^*)^2 \tag{3.2}$$

The monetary authority utility function is modeled with an analogous structure but with different preference weights, including the aversion to change its own instrument, the real interest rate (r), from its long-run level (equation 3.3). Let U^M represent the monetary authority's utility function and α_u^M, β_u^M and τ_u measure the cost associated to, respectively, an output fall ($\min(y - y^*, 0)$), an inflation increase ($\max(\pi - \pi^*, 0)$), and a real interest rate deviation ($r - r^*$). Note that α_u^M, β_u^M, $\tau_u > 0$. Again, we need to normalize the parameters, and for convenience this time let's do it with respect to the cost of output deviations: $\beta^M = \beta_u^M/\alpha_u^M$ and $\tau = \tau_u/\alpha_u^M$ (equation 3.4).

$$U^M = U^M\{(y - y^*), (\pi - \pi^*), (r - r^*)\}$$

[2] As will become clear below, working with asymmetric utility functions is important because it prevents the cancelling out of the policy biases over time: as shocks hit the economy, the fiscal and monetary authorities react by increasing the public deficit and the interest rate, respectively.

$$U^M = -\alpha_u^M\{\min(y - y^*, 0)\}^2 - \beta_u^M\{\max(\pi - \pi^*, 0)\}^2 - \tau_u(r - r^*)^2 \qquad (3.3)$$

$$U^M = -\{\min(y - y^*, 0)\}^2 - \beta^M\{\max(\pi - \pi^*, 0)\}^2 - \tau(r - r^*)^2 \qquad (3.4)$$

As pointed out above, the key assumption of the model lies in the difference between the monetary and fiscal authorities regarding their relative preferences for output and inflation deviations. We want our model to reflect both the Central Bank's mission to contain inflation and the fiscal authority's greater incentives to reduce unemployment. Thus, we assume that the monetary authority is more affected by inflation than output deviations. Conversely, the fiscal authority is more concerned about output than inflation deviations. In terms of the parameters of our model,

$$\alpha^F = \alpha_u^F/\beta_u^F > 1 \qquad \text{and} \qquad \beta^M = \beta_u^M/\alpha_u^M > 1$$

We assume that the long-run levels y^*,r^*, D^*, and r^* are perceived to be the same by both authorities.

The assumption that $\tau > 0$ reflects the fact that central banks dislike large and sudden movements in policy interest rates. The literature on monetary policy takes this into account by using the lagged interest rate as an argument of the policy reaction function (see Woodford 1999 for a model that formalizes the optimality of interest-rate smoothing rules.) Similarly, the assumption that $\delta > 0$ reflects the fact that fiscal authorities dislike deviations in their instrument from an established target. This may result from the political costs and financing difficulties involved in moving away from an agreed-upon fiscal budget.

The forces that rule the economy are modeled as follows:

$$y - y^* = \gamma_D(D - D^*) - \gamma_r(r - r^*) + \gamma_0 \qquad (3.5)$$

$$\pi - \pi^* = \lambda_y(y - y^*) - \lambda_0 \qquad (3.6)$$

Equation (3.5) represents the aggregate demand function and (3.6), the aggregate supply function (or Phillips Curve). The parameters γ_D and γ_r represent the elasticities of the output gap to the fiscal deficit and the real interest rate, respectively.

The parameter λ_y represents the elasticity of inflation to the output gap. Aggregate demand and supply shocks are represented by γ_0 and λ_0, respectively. For simplicity, we set $D^*, r^* = 0$. [3]

In what follows we concentrate on aggregate supply shocks. These are the most interesting given that for both authorities they elicit a trade-off between restoring output or controlling inflation. [4]

The solution for the case of a positive aggregate supply shock ($\lambda_0 > 0$) is trivial. This comes from the type of asymmetric loss functions that we assume. A positive supply shock leaves inflation lower than π^* and output higher than y^*, in which case neither authority suffer a loss and thus their is no policy response. On the other hand, a negative supply shock lowers output and increases inflation, thus inducing a policy reaction by both authorities. It is this case which we study in detail in the following sections.

Given a negative supply shock, the loss functions (equations 3.2 and 3.4) can be written as follows:

$$U^F = -\alpha^F(y - y^*)^2 - (\pi - \pi^*)^2 - \delta(D - D^*)^2 \tag{3.7}$$

$$U^M = -(y - y^*)^2 - \beta^M(\pi - \pi^*)^2 - \tau(r - r^*)^2 \tag{3.8}$$

This simplification is consistent with the resulting equilibrium level of output, which is lower than y^*, and of inflation, which is higher than π^*. In other words, the solution is interior to the range $y < y^*$ and $\pi > \pi^*$n (see below).

[3] A standard version of the modern Phillips Curve would look like $\pi = (y - y^*) + \pi^e - \lambda_0'$. If we subtract π^* from both sides we get $\pi - \pi^* = (y - y^*) - \lambda_0$ where $\lambda_0 = (\pi^* - \pi^e + \lambda_0')$. Our simple model does not provide a framework to generate endogenous expectations, so π^e is taken as an exogenous variable. Given π^*, we treat misaligned expectations with respect to π^* as supply shocks ($\pi^e > \pi$ would be a negative supply shock).

[4] It can be shown that even in the presences of a demand shocks the result that an increase in preference divergence leads to higher fiscal deficits and higher real interest rates holds. For a reference, see Bennett and Loayza 2000.

3.1.2 The Nash equilibrium

The relationship between fiscal and monetary authorities lies somewhere in between complete policy coordination and none at all [5]. The degree of coordination between fiscal and monetary authorities varies from country to country and over time. Our contention, however, is that a situation of lack of full coordination is the most prevalent when the monetary authority is independent from the central government. We arrive at this assumption considering the difficulties to achieve successful coordination, including the obstacles to enforce commitments between independent branches of government, the practical inability to discern outcomes due to policies from those due to exogenous shocks, and the transaction costs of day-to-day policy harmonization and monitoring. We model uncoordinated policy responses through a Nash game. The concept of equilibrium that the Nash equilibrium represents is an extreme case (no coordination at all). However it allows us to isolate the effect of uncoordinated actions that we want to analyze.

In our monetary-fiscal game, each authority decides the level of its instrument knowing that its counterpart is rational and has certain preferences over inflation and output gaps. The simple Nash equilibrium applies when both players decide simultaneously and without coordination their respective strategies. It consists of a pair (D^N, r^N) characterized by the property that no player can reach a higher level of utility by unilaterally deviating from this solution.

The Nash equilibrium is obtained when each authority maximizes its utility function with respect to its own instrument, taken the other policy instrument as given. Maximizing the fiscal authority's utility function (eq. 3.7), and substituting the expressions for aggregate demand and supply (eqs. 3.5 and 3.6), we obtain the fiscal reaction function (FRnFn):

$$\begin{aligned}\frac{\partial U^F}{\partial D} &= 0 \Longrightarrow \\ \text{FRnFn:} \quad D &= \left[\frac{1}{1+\frac{\delta}{\gamma_D^2(\alpha^F+\lambda_y^2)}}\right]\frac{\gamma_r}{\gamma_D}r + \left[\frac{\lambda_y\lambda_0}{\frac{\delta}{\gamma_D}+\gamma_D(\alpha^F+\lambda_y^2)}\right] \qquad (3.9)\end{aligned}$$

Similarly, maximizing the monetary authority's utility function (eq. 3.8), we ob-

[5]By coordination we understand the process through which two independent authorities negotiate their strategies in order to improve results for both.

tain the monetary reaction function (MRnFn):

$$
\begin{aligned}
\frac{\partial U^M}{\partial r} &= 0 \Longrightarrow \\
\text{MRnFn:} \quad D &= \left[1 + \frac{\tau}{\gamma_r^2(1+\beta^M\lambda_y^2)}\right]\frac{\gamma_r}{\gamma_D}r + \left[\frac{\lambda_y\lambda_0}{\gamma_D(1/\beta^M + \lambda_y^2)}\right]
\end{aligned}
\tag{3.10}
$$

Both reaction functions imply a linear relation between D and r. Comparing the two equations we can see the following. i) The intercept of MRnFn is more negative than the intercept of FRnFn. ii) The slope of MRnFn is higher than the slope of the iso-aggregate demand lines (γ_r/γ_D), and this in turn is higher than the slope of FRnFn. This results from the assumption that the monetary authority loses more from inflation than output gaps, and conversely for the fiscal authority, as well as, from the welfare loss to each authority because of deviations of their instrument from its long-run level. Establishing these differences between the two reaction functions will help us make the comparative static exercises presented below.

The intersection of MRnFn and FRnFn gives the Nash solution, as depicted in Figure 2. After a fair amount of algebra, the Nash equilibrium solution for the deficit and the interest rate is given by, [6]

$$
D^N = \frac{-\gamma_r^2\lambda_y(\alpha^F\beta^M - 1)\lambda_0 + \tau\lambda_y\lambda_0}{\gamma_r^2\delta/\gamma_D(1+\beta^M\lambda_y^2) + \tau\delta/\gamma_D + \gamma_D\tau(\alpha^F + \lambda_y^2)} \tag{3.11}
$$

$$
r^N = \frac{-\gamma_D^2\lambda_y(\alpha^F\beta^M - 1)\lambda_0 - \delta\beta^M\lambda_y\lambda_0}{\gamma_r\delta(1+\beta^M\lambda_y^2) + \tau\delta/\gamma_r + \gamma_D^2\tau/\gamma_r(\alpha^F + \lambda_y^2)} \tag{3.12}
$$

We are now prepared to derive the central result from the theoretical model. How will the deficit and interest rate change if the difference in relative preferences between the two authorities becomes larger (i.e., if either α^F or β^M increases)? The

[6]Substituting the resulting D^N and r^N in the output gap equation (3.5) and in the inflation equation (3.6), it can be seen that the equilibrium output is lower than y^* and inflation higher than π^*. In other words, the solution is interior to the range $y < y^*$ and $\pi > \pi^*$. Thus, it is valid to assume that given a negative supply shock the authorities' utility function can be modeled as the simple quadratic form of equations (3.7) and (3.8).

answer is that both the deficit and the interest rate increase, reflecting the relatively more extreme position taken by both authorities. This can be shown directly, though tediously, by taking partial derivatives of D^N and r^N with respect to α^F and β^M. More insight can be gained by deriving this result from shifts in the reaction functions.

Figure 3 illustrates the case when β^M increases: the intercept of MRnFN becomes more negative its slope declines (reflecting the central bank's lower desired aggregate demand level given its stronger anti-inflationary preference). As Figure 3 shows, the new equilibrium $(\widehat{D}^N, \widehat{r}^N)$ will necessarily be located to the north-east of (D^N, r^N), which means higher levels of both policy instruments. Likewise, if α^F rises, then the intercept of FRnFn becomes less negative and its slope increases. This also leads unambiguously to an increase in the deficit and the interest rate. [7]

3.1.3 The Stackelberg equilibrium

Whereas the Nash equilibrium is obtained when both players move simultaneously, a sequential play of the game leads to the Stackelberg equilibrium.[8] For the monetary-fiscal game this means that one authorities decides first the magnitude of its instrument and the other one follows it. We assume that the monetary authority is the leader. The opposite case yields similar conclusions.

[7] An increase in the deficit or the interest rate can be characterized as a "policy bias" only if it intensifies the departure from the long-run levels D^* and r^*, normalized to zero in the solution of the model. This occurs if the Nash equilibrium for a given negative supply shock is such that both D^N and r^N are positive. As can be seen from equation (3.12), this is always the case for the interest rate; the conceptual reason is that a decrease in the interest rate with respect to $r*$ would increase inflation further than just the mere effect of the negative supply shock. However, from equation (3.11) note that D^N may be negative, as this carries the possibility of making the output and inflation gaps share the burden of the supply shock. To rule out the possibility of a negative D^N, we need to impose the condition that both authorities have sufficiently strong relative preferences over output or inflation gaps. Specifically, we assume that the condition given in equation (3.13) holds.

$$\alpha^F \beta^M - 1 > \frac{\tau}{\gamma_r^2} \tag{3.13}$$

The implication of this condition is fairly intuitive and appealing: if condition (3.13) does not hold, it will imply that in the presence of a negative supply shock, the equilibrium level of inflation with an independence central bank will be higher than the resulting equilibrium level of inflation assuming a similar model but with the fiscal authority controlling both instruments.

[8] By Stackelberg equilibrium we refer only to the Nash equilibria of the sequential move-game associated with its backward-induction outcome.

The reaction of the follower in the Stackelberg game is the same as in the Nash game: the fiscal reaction function, (3.9), is the optimal response of the fiscal authority to a given interest rate decided by the leader. Thus, the follower's (fiscal) reaction function is given by,

$$\text{FRnFn:} \quad D = \left[\frac{1}{1+\frac{\delta}{\gamma_D^2(\alpha^F+\lambda_y^2)}}\right]\frac{\gamma_r}{\gamma_D}r + \left[\frac{\lambda_y\lambda_0}{\frac{\delta}{\gamma_D}+\gamma_D(\alpha^F+\lambda_y^2)}\right] \quad (3.9)$$

The central bank's F.O.C. as the leader of the Stackelberg game is obtained by maximizing U^M with respect to r, taking into account that the central bank is now able to affect D according to the fiscal authority's reaction function (3.9). We can then express the monetary authority's "action" function (MAnFn) as follows,

$$\text{MAnFn:} \quad D = \left[1+\frac{\tau}{\Phi\gamma_r^2(1+\beta^M\lambda_y^2)}\right]\frac{\gamma_r}{\gamma_D}r + \left[\frac{\lambda_y\lambda_0}{\gamma_D(1/\beta^M+\lambda_y^2)}\right] \tag{3.14}$$

where

$$\Phi = \frac{1}{1+\frac{\gamma_D^2(1+\beta^F\lambda_y^2)}{\delta}} < 1 \tag{3.15}$$

Substituting (3.9) into (3.14), we can determine the central bank's optimal interest rate, r^S. Then, given r^s the fiscal authority decides its deficit according to (3.9). The Stackelberg solution is given by, [9]

$$D^S = \frac{-\Phi\gamma_r^2\lambda_y(\alpha^F\beta^M-1)\lambda_0+\tau\lambda_y\lambda_0}{\Phi\gamma_r^2\delta/\gamma_D(1+\beta^M\lambda_y^2)+\tau\delta/\gamma_D+\gamma_D\tau(\alpha^F+\lambda_y^2)} \tag{3.16}$$

$$r^S = \frac{-\Phi\gamma_D^2\lambda_y(\alpha^F\beta^M-1)\lambda_0-\Phi\delta\beta^M\lambda_y\lambda_0}{\Phi\gamma_r\delta(1+\beta^M\lambda_y^2)+\tau\delta/\gamma_r+\gamma_D^2\tau/\gamma_r(\alpha^F+\lambda_y^2)} \tag{3.17}$$

How does the Stackelberg solution compare with the Nash solution? The Stackelberg equilibrium produces lower deficits and interest rates because the sequential nature of the game allows the leader to elicit a more moderate response from the follower. In fact, note that the presence of the parameter Φ in the monetary reaction function is equivalent to an increase in the τ, the willingness of the central bank to change its instrument. Figure 4 illustrates the comparison between the two types

[9] With the same procedure used in the Nash game, it can be shown that the Stackelberg solution is interior to the range $y < y^*$ and $\pi > \pi^*$. Thus, it is valid to assume that given a negative supply shock the authorities' utility function can be modeled as the simple quadratic form of equations (3.7) and (3.8).

of solutions. The monetary action function in the Stackelberg game has the same intercept, but a steeper slope than the monetary reaction function in the Nash game. Thus rendering a Stackelberg equilibrium with lower deficit and interest rate than the Nash equilibrium.

What happens in the Stackelberg game if the difference in relative preferences between the two authorities becomes larger? Since the preference parameters affect the monetary and fiscal reaction functions qualitatively the same in the Stackelberg as in the Nash games, an increase in either α^F or β^M produces larger deficits and interest rates. The only difference is that these changes in the policy instruments are attenuated because the sequential nature of the game acts as if signaling a lower willingness on the part of the leader to change its instrument. [10]

3.2 Empirical Evidence

The main result of the theoretical section can be presented as follows. When the central bank's anti-inflationary stance gets stronger or the fiscal authority becomes keener to avoid output gaps, then, in the face of a negative supply shock, the interest rate and public deficit will be larger as each authority takes a more extreme policy position to achieve its respective objective. In the empirical section we are interested in testing one important implication from this result. Namely, we want to test whether a change in the preference stance of one authority will produce a larger policy response in the other. Naturally, there are two components to this test. The first is whether public deficits will be larger when the monetary authority becomes more adamant in its anti-inflationary objective. The second is whether interest rates will be higher when the fiscal authority becomes more concerned about output gaps.

Our empirical approach will be based on cross-country and time-series comparisons. Then, we will be testing whether in countries and time periods where one authority becomes more extreme in its preferences, the other's policy instrument turns out to be stronger. In order to make valid cross-country and time-series com-

[10]The condition for the equilibrium fiscal deficit in the Stackelberg game to be positive is analogous to that in the Nash equilibrium (footnote 7). For the Stackelberg case we assume that this condition holds.

$$\alpha^F \beta^M - 1 > \frac{\tau}{\gamma_r^2 \Phi} \tag{3.18}$$

parisons, we need to normalize the data to make it consistent (as explained below) and, especially, we need to control for the size of the shocks and economic conditions that elicit government interventions in the first place. We will work under the maintained assumption that the fiscal and monetary reactions are not fully coordinated and respond to different relative preferences in the countries and time periods under consideration.

Before discussing the details of our empirical exercise, we should note that there is another implication of the model that we will not be focusing on. This is the implication that each authority's policy instrument responds to its own preferences. The reason why we will not use this implication for testing the model is that each authority's own preferences affect its instrument for a variety of reasons unrelated to the other authority's behavior. For instance, more fiscally responsible governments will reduce the deficit, independently of what the central bank's reaction might be. Thus, we could not claim a deficit's reduction in the case of a less activist government as supporting evidence for the model.

3.2.1 The empirical model

Let d be the primary deficit (properly normalized to be comparable across countries and over time) and r the real domestic interest rate (specifically, its portion subject to changes in policy.) Then, consider the following two reduced-form regression equations:

$$d_{i,t} = \beta_d X_{i,t}^d + \delta_d fp_{i,t} + \theta_d mp_{i,t} + \eta_i + \varepsilon_{i,t} \tag{3.19}$$

$$r_{i,t} = \beta_r X_{i,t}^r + \delta_r fp_{i,t} + \theta_r mp_{i,t} + \mu_i + \epsilon_{i,t} \tag{3.20}$$

where *fp* is an indicator of the fiscal authority's preference regarding its willingness to avoid output gaps, *mp* is an indicator of the monetary authority's preference on the importance of avoiding inflation, X is a set of control variables that includes a time trend, η and μ are unobserved country-specific effects, and the subscripts i and t denote country and year, respectively. We assume that there is cross-country and time-series homogeneity regarding all slope parameters.

Hypothesis test. The test of our main hypothesis is based on the sign and significance of both θ_d and δ_r. If both are significantly positive, we conclude that an increase in the preference divergence between monetary and fiscal authorities leads to an increase in their respective policy instruments, *ceteris paribus.*

Sample. We use a pooled data set of annual observations for the period 1970-94 covering most industrialized countries. Since the paper focuses on the interaction between fiscal and monetary policies towards stabilization, we cannot use countries where this relationship has been dominated by the inflationary finance of the fiscal deficit. We recognize that in that case, the issues analyzed in this paper are of second order importance. Therefore, in the empirical analysis we do not work with developing countries or with OECD countries that have experienced relatively high inflation over the last three decades. For this reason we have not included countries whose average inflation rate for the period has been above 10%]; this is the case of Greece (average inflation: 14%), Iceland (24%), Italy (10.5%), and Portugal (14%.) For the remaining industrial countries the average inflation rate for the period has been below 10%. The countries in the sample are Australia, Austria, Belgium, Canada, Switzerland, Germany, Denmark, Spain, Finland, France, the United Kingdom, Ireland, Japan, Netherlands, Norway, New Zealand, Sweden, and the United States. [11]

Dependent variables. For the deficit regression, the dependent variable is the ratio of the central government's primary deficit (total deficit minus interest payments) to GDP (obtained from *Government Finance Statistics*, IMF). Dividing by GDP normalizes the scale (or metric) of primary deficits such that it can be used for regressions across countries and over time. For the interest-rate regression, the dependent variable is the difference between domestic and international real interest rates. We use the deviations from the international rate to account for the fact that domestic rates are heavily influenced by parity conditions in countries with open capital accounts. Only the portion of domestic rates not determined by international conditions can be regarded as the monetary authority's policy instrument. Real domestic interest rates are obtained by deflating nominal domestic money-market rates (13-week Treasury Bill Rate or its equivalent) by an average of current and next-year CPI inflation rates (from Loayza *et al.* (1998)). The international real interest rate is the 13-week nominal Eurodollar London rate adjusted accordingly with the industrial countries average CPI (from *International Financial Statistics*, IMF).

Indicators for the preferences of fiscal and monetary authorities. We proxy for the fiscal authority's preference to avoiding output gaps with a measure of the ruling government's political orientation. Specifically, we use the Beck *et al.* (2000) political institutions data base to assess the political orientation of the Chief Executive's Party (*right*, *center* or *left*). We assume that parties on the left of the political spectrum have a higher preference for reducing unemployment and the output gap than controlling inflation, and *vice versa.* We constructed the indicator *fp* by assigning a value of 0 to countries and years where the political orientation of the Chief Executive's Party was defined as *right*, a value of 0.5 where it was defined as *center*, and a value of 1 where it was defined as *left*. Beck's *et al.* dataset covers 177 countries over 21 years, 1975-1995. For the period 1970-1974, we constructed the indicator using information from Banks and Muller (1998).

We proxy for the monetary authority's preference against inflation hikes by the formal central bank's commitment to control inflation as expressed in its charter. This indicator is obtained from the Cukierman, Webb and Neyapti (1992) cross-country database on central bank independence and objectives. Their database contains information on most industrialized countries and many developing countries over the period 1970-89, and we update it from central bank documents up to 1994. The measure of interest for our purposes is the Cukierman-Webb-Neyapti index for the importance of price stability as a central bank objective. They construct the index based on explicit information contained in each central banks' charter and assign specific values according to the following criteria: 1 if *price stability is mentioned as the only or major goal, and in case of conflict with the government, the central bank has final authority to pursue policies aimed at achieving this goal*; 0.8 if *price stability is mentioned as the only goal*; 0.6 if *price stability is mentioned along with other objectives that do not seem to conflict with price stability (e.g., stable banking)*; 0.4 if *price stability is mentioned with a number of potentially conflicting goals (e.g., full employment)*; 0.2 if *central bank charter does not contain any objectives for the central bank* and 0 if *some goals appear in the charter but stability is not one of them.* In a further exercise we also use the Cukierman-Webb-Neyapti index of central bank independence, (cbi), exclusive of the price-stability objective, in order to compare the effects of central bank independence with those of central bank preference for price stability.

We work under the notion that the preferences of monetary and fiscal authorities are mainly driven by institutional and political factors and possibly as well by economic conditions. However, we assume that these preferences are exogenous with respect to the authorities' current policy stance. In other words, we assume that there is no reverse causation from public deficits and interest rates to the authorities' preferences.

Control and instrumental variables. Given that our analysis is based on the comparison across countries and time periods, we need to control for the economic conditions and shocks that prompt a monetary or fiscal reaction in the first place. As summary measures for these conditions and shocks, we include the rate of GDP growth per capita, the rate of CPI inflation, and the rate of saving as explanatory variables [12]. It is likely that these variables be endogenous with respect to the policy instruments, our dependent variables. This would not be problem in testing the hypothesis of interest if growth, inflation, and saving were uncorrelated with the authorities' preferences. However, this may not necessary be the case, and there could be a correlation between these preferences and economic conditions. Therefore, in some of the exercises presented below we use an instrumental-variable procedure to isolate the exogenous components of growth, inflation, and saving. As instruments, we use the trade-weighted average GDP growth of partner countries, the terms of trade shocks, the level and squared of (the log of) per capita GDP, the share of the population living in urban areas, the share of the population between the ages of 15 and 65, and the share of the population older than 65 years. The first two instrumental variables are mostly related to growth and inflation; the last three are mostly related to saving; and the instruments based on per capita GDP are related to all three explanatory variables. Data for the control and instrumental variables are obtained from the *World Development Indicators*, World Bank (various years), and Loayza *et.al* (1998).

As explained above, our test is based on the effect that the preferences of one of the authorities may have on the other's policy instrument. This effect, however, is conditional on the relative preferences of the latter. For instance, the effect that the central bank's inflation aversion may have on the public deficit is conditional on how averse the fiscal authority is with respect to unemployment. Therefore, although we are concerned with the coefficient of only one of the authorities' preferences in each regression, we must include both preference proxies as explanatory variables. Finally, all regressions control for a common time trend, and, in most specifications, they also control for unobserved country-specific effects.

3.2.2 Estimation and results

Table 1 is organized in three pairs of results, where each pair consists of a deficit and an interest-rate regression. The first pair is estimated using a simple pooled OLS methodology, where each observation is treated independently of which country it belongs to. The second pair of regressions is estimated with a fixed-effects estimator,

[12]National saving in the interest-rate regression, and private saving in the deficit regression.

whereby unobserved country-specific effects are captured by country intercepts [13]. For the third pair of regressions we still use the fixed-effects method but allow for a richer structure of the error term; specifically, using the Newey-West procedure, we correct the standard errors of the coefficients by allowing for country-wise heteroskedasticity and MA(1) residual correlation. All results support the hypotheses implied from the theoretical model, in the sense that the coefficients of interest are positive and statistically significant, as expected. That is, an increase in the anti-inflation preference of the central bank induces a higher primary public deficit; and, likewise, an increase in the anti-unemployment preference of the fiscal authority invites a higher real interest rate. The size of the coefficient on the fiscal authority preference index remains approximately the same whether we consider or not country fixed effects. Conversely, the coefficient on the monetary authority preference index does increase notably when country-specific effects are accounted for.

Could the effects that we attribute to central bank preferences be due in fact to its independence per se? To address this alternative explanation we re-run the fixed-effects estimator with robust standard errors adding to the regression the index for central bank independence, or alternatively, substituting it for the index on central bank preferences (see Table 2). We find that including the index of central bank independence does not drive out the positive impact of inflation aversion on the primary deficit. Moreover, central bank independence has no beating on the primary public deficit. Interestingly, central bank independence has a negative effect on the real interest rate, just the opposite of central bank's inflation aversion.

The last issue we examine is whether the potential endogeneity of the control variables may be spuriously driving the results. Table 3 shows the instrumental-variable estimation results, controlling in addition for unobserved country specific effects. The coefficient on the central bank preference index is positive and significant in the deficit regression, and so is the coefficient on the fiscal authority preference index in the interest-rate regressions. The size of both coefficients rise by about fifty percent once the endogeneity of the control variables is accounted for.

Finally, using the point estimates obtained in the last set of regressions, we can assess the economic importance of these effects. The estimated coefficients with the instrumental-variable estimation imply that a 1 standard deviation increase in mp is associated with a primary deficit increase of 2 percent of GDP. As well, a 1 standard

[13]We had the option of modeling the unobserved effects as random, but using a Hausman-type test we rejected the hypothesis that the effects were uncorrelated with the explanatory variables.

deviation increase in fp is associated with a rise of the domestic real interest rate by 0.8 percentage points (with respect to the international rate).

3.3 Conclusions

Central bank independence has contributed to achieve price stability and fiscal discipline for many countries. The conventional wisdom is that this is a necessary, first-generation reform for not only monetary but also fiscal policy. The policy question this paper raises is whether a second-generation reform consisting of institutional incentives for domestic policy coordination could have potential benefits. The paper presents a game-theoretic model where the fiscal and monetary authorities interact to stabilize the economy in the face of negative supply shocks, having dissimilar preferences with respect to output and inflation gaps and controlling different policy instruments. It is assumed, realistically, that the monetary authority has a larger utility loss from inflation than output gaps and conversely for the fiscal authority.

Modeled as either Nash or Stackelberg equilibria, the solution under lack of policy coordination implies that, given a negative supply shock, the fiscal authority acts more liberally and the monetary authority more conservatively. In particular, we find that an increase in the preference divergence between the monetary and fiscal authorities leads to, *ceteris paribus*, larger public deficits (the fiscal authority's policy instrument) and real interest rates (the central bank's instrument).

The empirical section attempts to test an important implication of the model on a pooled sample of 18 industrial countries with annual information for the period 1970-94. In particular, we test whether in countries and time periods where one authority becomes more extreme in its preferences, the other takes a more radical position regarding its policy instrument in order to pursue its objective. We proxy for the fiscal authority's preference to avoiding unemployment and output gaps by the political orientation of the party in executive power, with leftist governments receiving a higher proxy value. We measure the monetary authority's preference against inflation hikes by the formal central bank's commitment to price stability, as expressed in its charter. Controlling for economic conditions and shocks, we find that an increase in the anti-inflation preference of the central bank induces a higher primary public deficit. Likewise, we find that an increase in the anti-unemployment preference of the fiscal authority invites a higher difference of the domestic interest rate with respect to the international rate. These results hold true when we control for economic conditions and shocks affecting the countries in our sample. Moreover, they are statistically significant whether or not we allow for country-specific effects, common time trends, robust standard errors, and the possibility of endogenous control variables. We believe these results provide empirical support to the hypothesis that under lack of coordination, the more divergent fiscal and monetary authorities are in their preferences, the more radical positions they will take, potentially hurting the economy along the way.

The main policy implication from the paper is that, notwithstanding the gains from central bank independence, coordination both at the level of setting objectives and at the level of policy implementation can alleviate the biases that move the economy to sub-optimally higher fiscal deficits and interest rates. This goal can be achieved with "second generation reforms" that should deal with the difficulties of policy coordination, such as contract enforceability and the practical inability to discern outcomes due to policies from those due to shocks.

3.A Tables and Figures

Table 1: The Effects of Different Preferences Between Fiscal and Monetary Authorities on Fiscal Deficits and Real Interest Rates: Core Specification

Sample: 18 Industrial Countries, 1970 – 1994

Estimation: (a) Pool OLS (b) FE (c) Robust FE

Dependent Variable		Pool OLS (a) D/GDP	Pool OLS (a) $r\text{-}r^{int}$	FE (b) D/GDP	FE (b) $r\text{-}r^{int}$	Robust FE (c) D/GDP	Robust FE (c) $r\text{-}r^{int}$
Authorities'	***fp***	0.005	0.011	0.006	0.010	0.006	0.010
		(1.72)	(3.86)**	(1.95)	(3.18)**	(1.74)	(2.98)**
preferences	***mp***	0.011	0.009	0.039	0.019	0.039	0.019
		(2.44)**	(2.01)	(3.88)**	(1.82)	(2.52)**	(1.75)
Growth		-0.002	-0.004	-0.004	-0.004	-0.004	-0.004
		(-2.97)	(-5.86)	(-5.77)	(-5.96)	(-3.94)	(-6.06)
Inflation		0.125	-0.388	0.026	-0.433	0.026	-0.433
		(2.84)	(-9.05)	(0.52)	(-8.72)	(0.35)	(-7.74)
Private saving		0.263		0.378		0.378	
		(6.80)		(5.98)		(4.01)	
National saving			-0.010		0.106		0.106
			(-0.32)		(1.67)		(1.55)
Trend		-0.001	0.001	-0.001	0.001	-0.001	0.001
		(-1.97)	(3.81)	(-4.29)	(3.29)	(-3.52)	(2.76)
R^2		0.17	0.42	0.21	0.42	0.21	0.42
#Countries / Obs.		18/421	18/421	18/421	18/421	18/421	18/421

Notes: (a) (b) t-Statistics in parenthesis, (c) NW robust t-Statistics in parenthesis. Only for the cross-preference variable (highlighted in gray): ** significant at 5%, * significant at 10%.

Table 2: The Effects of Different Preferences Between Fiscal and Monetary Authorities on Fiscal Deficits and Real Interest Rates: Central banks' preference for inflation and independence

Sample: 18 Industrial Countries, 1970 – 1989

Estimation: Robust FE

		Robust FE			Robust FE		
Dependent Variable		D/GDP	D/GDP	D/GDP	r-r^{int}	r-r^{int}	r-r^{int}
Authorities'	***fp***	0.005	0.005	0.007	0.009	0.009	0.013
		(1.20)	(1.22)	(1.82)	(2.30)**	(2.47)**	(3.31)**
preferences	***mp***	0.058	0.063		0.064	0.091	
		(3.96)**	(3.99)**		(2.76)	(3.78)	
	cbi		-0.053	0.016		-0.276	-0.175
			(-0.81)	(0.20)		(-4.06)	(-2.39)
Growth		-0.003	-0.003	-0.003	-0.002	-0.002	-0.002
		(-3.98)	(-3.92)	(-3.96)	(-2.21)	(-2.40)	(-2.37)
Inflation		0.046	0.044	0.036	-0.341	-0.355	-0.365
		(0.78)	(0.75)	(0.59)	(-5.68)	(-6.15)	(-5.98)
Private saving		0.294	0.286	0.284			
		(3.01)	(2.91)	(2.81)			
National saving					-0.078	-0.075	-0.091
					(-0.91)	(-0.89)	(-1.05)
Trend		-0.001	-0.001	-0.001	-0.000	-0.000	-0.000
		(-2.85)	(-2.94)	(-2.45)	(-0.63)	(-0.76)	(-0.35)
R^2		0.21	0.18	0.21	0.25	0.21	0.22
#Countries / Obs.		18/336	18/336	18/336	18/336	18/336	18/336

Notes: NW robust t-Statistics in parenthesis. Only for the cross-preference variable (highlighted in gray): ** significant at 5%, * significant at 10%.

Table 3: The Effects of Different Preferences Between Fiscal and Monetary Authorities on Fiscal Deficits and Real Interest Rates: IV Estimation

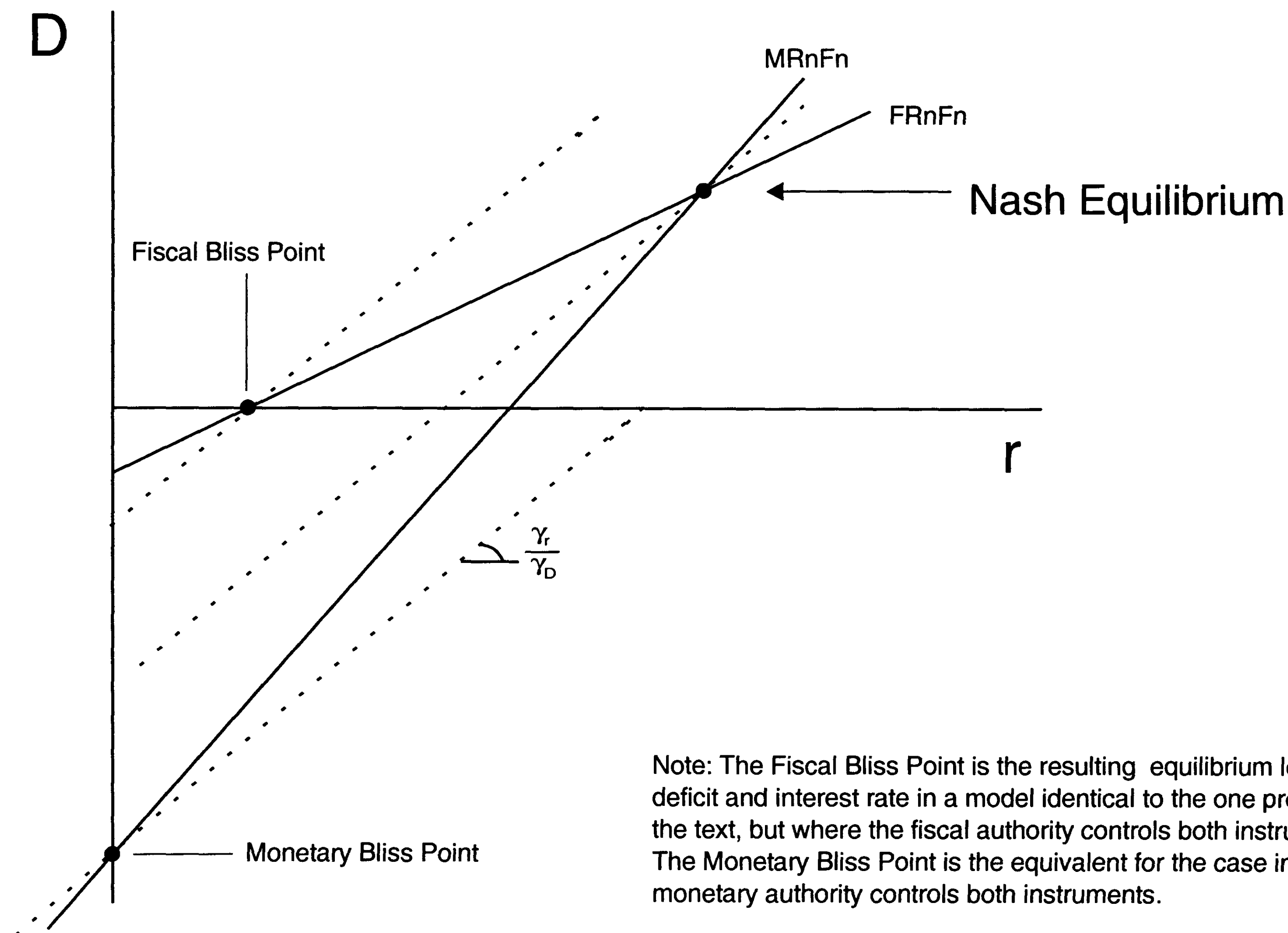
Sample: 18 Industrial Countries, 1970 – 1994

Estimation: (a) IV-FE (b) Robust IV-FE

		IV FE (a)		Robust IV FE (b)	
Dependent Variable		D/GDP	r-r^{int}	D/GDP	r-r^{int}
Authorities'	***fp***	0.000	0.017	0.000	0.017
		(0.02)	(2.83)**	(0.01)	(2.25)**
preferences	***mp***	0.060	0.004	0.060	0.004
		(2.52)**	(0.19)	(2.11)**	(0.15)
Growth		-0.003	-0.014	-0.003	-0.014
		(0.86)	(-3.47)	(0.78)	(-3.09)
Inflation		0.623	-1.035	0.623	-1.035
		(1.34)	(-2.47)	(1.15)	(-1.99)
Private saving		0.188		0.188	
		(0.49)		(0.48)	
National saving			0.251		0.251
			(1.33)		(1.12)
Trend		0.001	-0.001	0.001	-0.001
		(0.40)	(0.82)	(0.37)	(0.64)
R^2		0.17	0.35	0.17	0.35
#Countries / Obs.		18/421	18/421	18/421	18/421

Notes: (a) t-Statistics in parenthesis, (b) NW robust t-Statistics in parenthesis. Only for the cross-preference variable (highlighted in gray): ** significant at 5%, * significant at 10%. Exogenous variables: ***fp*** and ***mp***; Exogenous instruments: Foreign growth rate weighted by trade share, terms of trade, logarithm of GDP p/c, logarithm of GDP p/c squared, urban population (% of total population), active population 15-65 (% of total population), senior population 65+ (% of total population).

Figure 2. The Nash equilibrium



Note: The Fiscal Bliss Point is the resulting equilibrium level of deficit and interest rate in a model identical to the one presented in the text, but where the fiscal authority controls both instruments. The Monetary Bliss Point is the equivalent for the case in which the monetary authority controls both instruments.

Figure 3. An increase in central bank's anti-inflation preference (Nash equilibrium)

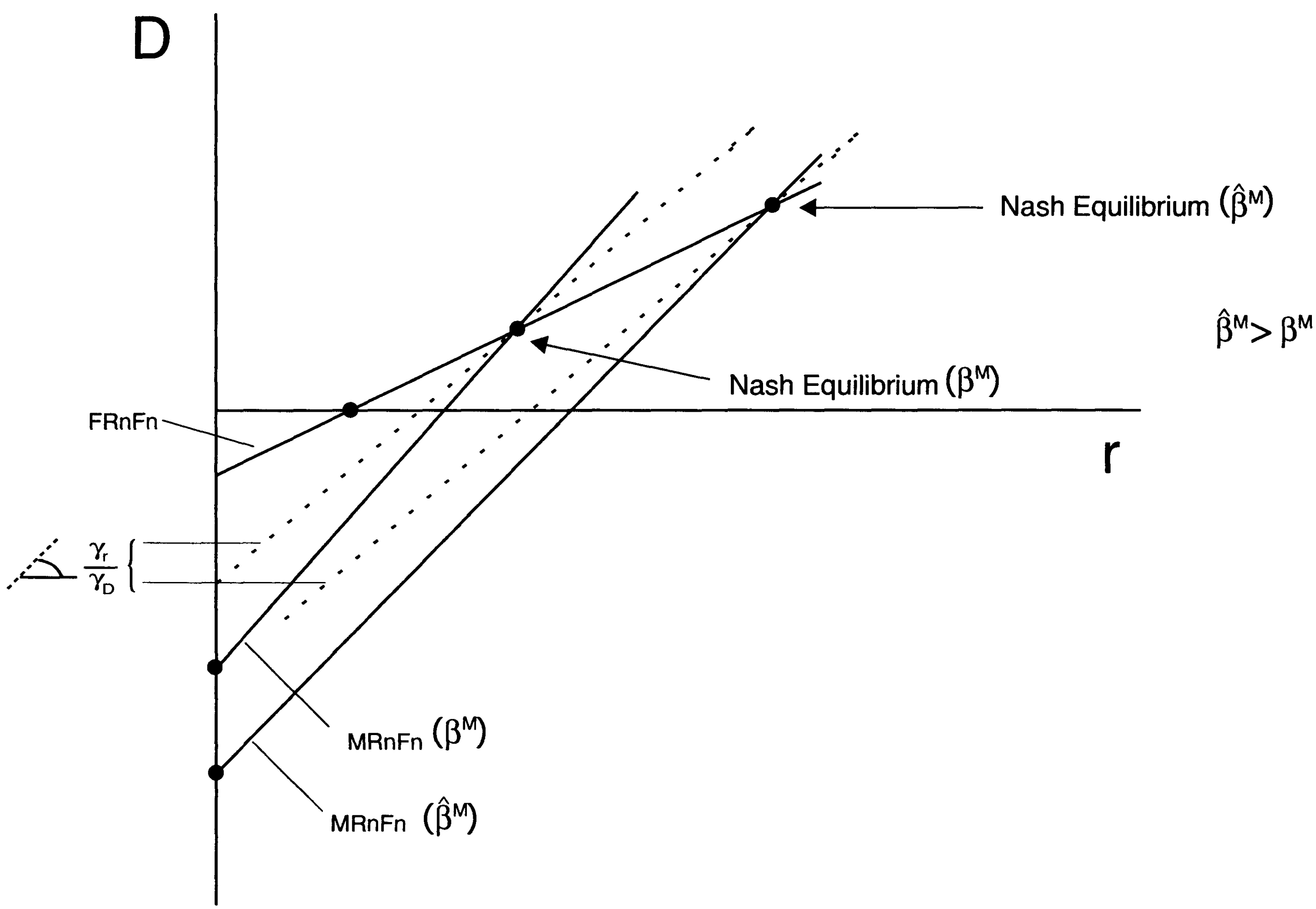
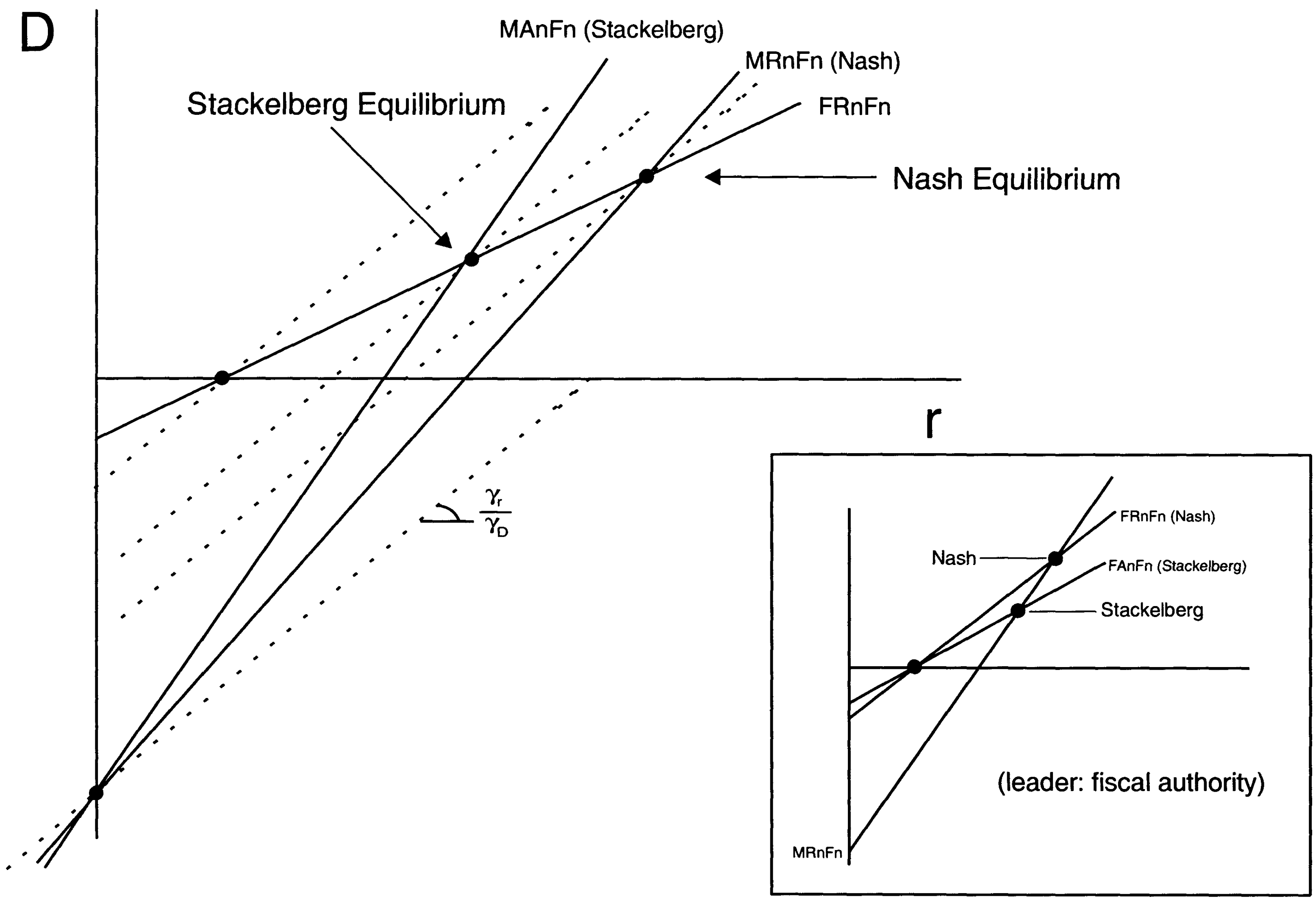


Figure 4. The Stackelberg equilibrium
(leader: central bank)



Bibliography

[1] Acemoglu, D. and F. Zilibotti (1997). "Was Prometheus Unbound by Chance? Risk, Diversification and Growth." *Journal of Political Economy* 105: 709–751.

[2] Alvarez, F. and M. Veracierto (2001). "Severance Payments in an Economy with Frictions." *Journal of Monetary Economics* 47(3): 477–498.

[3] Bai, J. and P. Perron (1998). "Estimating and Testing Linear Models with Multiple Structural Changes." *Econometrica* 66(1): 47-78.

[4] Bai, J. and P. Perron (2003). "Computation and Analysis of Multiple Structural Change Models." *Journal of Applied Econometrics* 18: 1-22.

[5] Banks, A. and T. Muller (1998). *Political Handbook of the World, 1998 (Serial).* CSA Publications, Binghamton University, New York.

[6] Beck, T., G. Clarke, A. Groff, P. Keefer, and P. Walsh (2000). "New Tools and New Tests in Comparative Political Economy: The Database of Political Institutions." *World Bank Policy Research Working Paper* 2283, February.

[7] Bennett, H., and N. Loayza (2000). "Policy Biases when the Monetary and Fiscal Authorities have Different Objectives." *Working Paper Series of the Central Bank of Chile* 66.

[8] Bentolila,S. and G. Bertola (1990). "Firing Costs and Labor Demand: How Bad is Eurosclerosis?" *Review of Economic Studies* 57: 381–402.

[9] Bernanke, B. S., T. Laubach, F. Mishkin, and A. Posen (1999). *Inflation Targeting. Lessons from the International Experience.* Princeton University Press.

[10] Blanchard, O. (2000). *The Economics of Unemployment. Shocks, Institutions, and Interactions.* Lionel Robbins Lectures, London School of Economics.

[11] Blanchard, O. and J. Simon (2001). "The Long and Large Decline in U.S. Output Volatility." *Brookings Papers on Economic Activity* 1: 135–164.

[12] Braun, M. (2003). "Financial Contractibility and Assets' Hardness." *Mimeo*, UCLA.

[13] Caballero, R. and M. Hammour (1998). "Jobless Growth: Appropriability, Factor Substitution and Unemployment." *Carnegie-Rochester Conference Series on Public Policy* 48: 51–94.

[14] Campbell, J., M. Lettau, B. Malkiel, and Y. Xu (2001). "Have Individual Stocks Become More Volatile? An Empirical Exploration of Idiosyncratic Risk." *Journal of Finance* 56(1): 1–43.

[15] Comin, D. and S. Mulani (2004). "Diverging Trends in Macro and Micro Volatility: Facts." *NBER Working Papers* 10922.

[16] Comin, D. and S. Mulani (2005). "A Theory of Growth and Volatility at the Aggregate and Firm Level." *NBER Working Papers* 11503.

[17] Comin, D. and T. Philippon (2005). "The Rise in Firm-Level Volatility: Causes and Consequences." *NBER Working Papers* 11388.

[18] Cukierman, A. (1992). *Central Bank Strategy, Credibility, and Independence: Theory and Evidence.* MIT Press.

[19] Cukierman, A., S. B. Webb and B. Neyapti (1992). "Measuring the Independence of Central Banks and Its Effect on Policy Outcomes." *The World Bank Economic Review* 6(3): 353-398.

[20] Davis S., J. Haltiwanger, S. Schuh (1996). "Job Creation and Destruction." MIT Press, Cambridge, Massachusetts.

[21] Davis, S., J. Haltiwanger, R. Jarmin, and J. Miranda (2006). " Volatility and Dispersion in Business Growth Rates: Publicly Traded versus Privately Held Firms." Presented at the NBER's Twenty-first Annual Conference on Macroeconomics on April 7, 2006.

[22] Diamond, D. and G. Rajan (2001). "Liquidity Risk, Liquidity Creation, and Financial Fragility: A theory of Banking." *Journal of Political Economy* 109(2): 287–327.

[23] Frankel, J. A. (1988). "The Implications of Conflicting Models for Coordination between Monetary and Fiscal Policymakers." In *Empirical Macroeconomics for Interdependent Economies*, edited by Bryant, R.C. *et al.*. Washington, Brooking Institution.

[24] Gitman, L. (1974). "Corporate Liquidity Requirements: A Simplified Approach." *The Financial Review* 9: 79–88.

[25] Gilchrist, S. and M. Gertler (1994). "Monetary Policy, Business Cycles, and the Behavior of Small Manufacturing Firms." *Quarterly Journal of Economics.*

[26] Hart, O. and J. Moore (1994). "A Theory of Debt Based on the Inalienability of Human Capital." *Quarterly Journal of Economics* 109(4): 841–879.

[27] Hopenhayn, H. and R. Rogerson (1993). "Job Turnover and Policy Evaluation: A General Equilibrium Analysis." *Journal of Political Economy* 101(5): 915–938.

[28] Johansen, S. (1988). "Statistical Analysis of Cointegration Vectors." *Journal of Economic Dynamics and Control* 12: 231-254.

[29] Johansen, S. and K. Juselius (1990). "Maximum Likelihood Estimation and Inference on Cointegration–With Applications to the Demand for Money," *Oxford Bulletin of Economics and Statistics* 52(2): 169-210.

[30] Kim, C., and C. Nelson (1999). "Has the U.S. Economy Become More Stable? A Bayesian Approach Based on a Markov-Switching Model of the Business Cycle." *Review of Economics and Statistics* 81: 608–616.

[31] Knack, S. and P. Keefer (1995). "Institutions and Economic Performance: Cross-Country Tests Using Alternative Institutional Measures." *Economics and Politics* 7(3): 207–227.

[32] Lamont, O. (1997). "Cash flow and investment: Evidence from internal capital markets." *Journal of Finance* 52: 83–109.

[33] Lazear, E. (1990). "Job Security Provisions and Employment." *Quarterly Journal of Economics* 105(3): 699–726.

[34] Loayza, N., H. López, K. Schmidt-Hebbel, and L. Servén (1998). "The World Saving Database." *World Bank manuscript*, The World Bank, Washington, DC.

[35] Loayza, N. and R. Ranciere (2006). "Financial Development, Financial Fragility, and Growth." Forthcoming *Journal of Money, Credit, and Banking.*

[36] Loewy, M. (1988). "Reganomics and Reputation Revisited." *Economic Inquiry.* Vol. XXVI (April): 253-263.

[37] McConnell, M. and G. Perez-Quiros (2000). "Output Fluctuations in the United States: What Has Changed Since the Early 1980s?" *American Economic Review* 90(5): 1464–76.

[38] Nordhaus, W. (1994). "Policy Games: Coordination and Independence in Monetary and Fiscal Policies." *Brooking Papers on Economic Activity* 2: 139-215.

[39] Obstfeld, M. (1994). "Risk-Taking, Global Diversification, and Growth." *American Economic Review* 84: 1310–1329.

[40] OECD (2004). "Employment Outlook." Organisation for Economic Co-Operation and Development, Paris, France.

[41] Perry, G. and D. Lederman (1997). "The Importance of Macroeconomic Policy Coordination: A Note". *Mimeo*, Office of the Chief Economist, Latin America and the Caribbean, The World Bank.

[42] Pesaran, H. (1997). "The Role of Econometric Theory in Modeling the Long Run." *Economic Journal* 107: 178-191.

[43] Pesaran, H. and Y. Shin (1999). "An Autoregressive Distributed Lag Modelling Approach to Cointegration." Chapter 11 in *Econometrics and Economic Theory in the 20th Century: The Ragnar Frisch Centennial Symposium*, Cambridge University Press.

[44] Philippon, T. (2003). "An Explanation for the Joint Evolution of Firm and Aggregate Volatility." NYU mimeo.

[45] Raddatz, C. (2005). "Liquidity Needs and Vulnerability to financial underdevelopment." *Journal of Financial Economics*, forthcoming.

[46] Rajan, R. and L. Zingales (1998). "Financial Dependence and Growth." *American Economic Review* 88(3): 559–586.

[47] Rauh, J. (2005). "Investment and Financing Constraints: Evidence from the Funding of Corporate Pension Plans." *Journal of Finance*, forthcoming.

[48] Stock, J. and M. Watson (2003). "Has The Business Cycle Changed And Why?" *NBER Macro Annual* 2003.

[49] Veracierto, M. (2001). "Employment Flows, Capital Mobility, and Policy Analysis." *International Economic Review* 42(3): 571–595.

[50] Walsh, C. E. (1993). "Central Bank Strategies, Credibility and Independence: A Review Essay." *Journal of Monetary Economics* 32: 287-302.

[51] Woodford, M (1999). "Optimal Monetary Policy Inertia." *NBER Working Paper* No. 7261, August.