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EQUALITY UNDER THREAT BY THE TALENTED: EVIDENCE FROM WORKER-MANAGED FIRMS*

Gabriel Burdín

Does workplace democracy engender greater pay equality? Are high-ability individuals more likely to quit egalitarian organisational regimes? The article revisits this long-standing issue by analysing the interplay between compensation structure and quit behaviour in the distinct yet underexplored institutional setting of worker-managed firms. The analysis is based on novel administrative data sources, which allow constructing a simple ordinal measure of the workers' ability type. The article's key findings are that worker-managed firms have a more compressed compensation structure than conventional firms, and high-ability members are more likely than other members to exit.

The potential conflict between equality and the need for incentives is a major debate in economics and political philosophy. Does workplace democracy engender greater pay equality? Are high-ability individuals more likely to quit from egalitarian regimes? I revisit this debate by analysing the relationship between compensation structure and quit behaviour in a unique and underexplored institutional setting: worker-managed firms (WMFs).

Most economic activities in actual market economies are carried out by conventional firms (CFs) controlled by capital suppliers. In contrast, WMFs are defined as enterprises in which the workforce has ultimate control rights (Dow, 2003). Their members have equal political influence on economic decisions regardless of their capital contribution to the firm ('one person, one vote'). This type of firm captured the attention of renowned economists such as Karl Marx, John Stuart Mill, Leon Walras, and Alfred Marshall. Since the late 1950s, an extensive theoretical literature has developed that seeks to understand the behaviour of WMFs and to explain why they are relatively rare.[1] One prominent explanation for the paucity of WMFs is that workplace democracy may result in substantial redistribution at the expense of high-ability workers.[2] Median voter models suggest that, to the extent the median member is less productive than the average, most cooperative members can gain by reducing wage differences relative to differences in productivity (Kremer, 1997). Another explanation

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[1] For a review of the literature see Bonin *et al.* (1993), Dow and Putterman (2000), Dow (2003) and Putterman (2008). The most updated evaluation of the empirical literature is provided by Pencavel (2013).

[2] Seminal theoretical analyses of how distribution rules affect WMFs include Sen (1966) and Gui (1987). See also Hansmann (1988).

is that equality may provide insurance against unfavourable realisations of ability (Abramitzky, 2008). Both models predict that equality discourages the participation of high-ability members. However, the actual extent and effects of redistribution in WMFs have not been systematically studied.

This article contributes to filling this gap by examining three interrelated questions. Do WMFs actually exhibit a more compressed compensation structure? Are high-ability members in WMFs more likely (than other members) to exit? Does the degree of equality affect the severity of brain drain? The empirical analysis is based on work history data from Uruguayan social security administrative records. To answer the first question, I use a panel of workers employed in both worker-managed and conventional firms. To address the second and third questions, I use a matched employer–employee panel data set that includes information on the total population of firms legally registered as producer cooperatives (PCs) – from which WMFs can be identified – and all their workers, both members and non-members. One major advantage of the latter data set is that I can observe the entire wage distribution at each firm for any given moment in time. This makes it possible to rank the ability of workers, including quitters, according to their position in the intra-firm wage distribution.

The analysis yields two main results. First, I find a small wage premium associated with being employed in WMFs. Because there is mobility between worker-managed and conventional firms, identification rests on the variability provided by workers who switch between organisational types during the period – under the assumption that sorting is based on time-invariant characteristics. It is noteworthy that this wage gap decreases across the wage distribution. Both pooled and fixed-effects quantile regression estimates confirm that WMFs exhibit a more compressed compensation structure than conventional firms. This result is consistent with the hypothesis that WMFs redistribute in favour of low-wage workers. Second, estimates derived from duration models indicate that the high-ability members of WMFs exhibit a higher hazard rate of voluntary separation.

At a more general level, this article contributes to the study of the interplay between equality and incentives that permeates many debates in public finance, development, comparative economic systems, human resources, and organisational economics. First, it is related to a series of recent studies on equal-sharing rules and migration in communes, particularly in Israeli kibbutzim (Abramitzky, 2008, 2009, 2011). The article adds to this literature in several ways. Kibbutzim studies have relied on self-reported measures of the degree of internal equality and have tested brain drain by comparing quitters to stayers in terms of education and skill levels, not in terms of their wages. Moreover, they have not investigated whether kibbutzim that shift away from equal-sharing rules do in fact reduce their brain drain. By contrast, I use matched organisation–worker panel data that give the entire wage distribution of each WMF and exploit within-firm variation in intra-firm wage dispersion to analyse how organisations use compensation policies to cope with brain drain. The interest in worker-managed firms rests on the fact that these organisations have existed (alongside investor-controlled firms) in most Western economies since the Industrial Revolution. Yet even though WMFs are thus a realistic organisational alternative to capitalist firms, they are usually found only in certain sectors (e.g. professional partnerships, taxis) and regions. The paucity of

WMFs, especially in labour-intensive sectors, remains a puzzle. Second, the choice of a compensation structure and its effect on the retention of valuable employees is a core topic in personnel economics (Lazear and Shaw, 2007; Lazear and Oyer, 2013). Third, the article is also related to the public economics literature on how mobility constrains redistributive taxation (Simula and Trannoy, 2010; Kleven *et al.*, 2013; Rothschild and Scheuer, 2013). This case study on WMFs illustrates how egalitarian schemes are threatened when some individuals have attractive exit options and so can 'vote with their feet'. Finally, this article contributes directly to the literature on WMFs by studying how members' heterogeneity and democratic governance actually interact in such firms (Pencavel, 2013). This study is one of the first to assess the extent and effects of redistributive compensation policies in WMFs.[3]

The rest of the article is organised as follows. Section 1 provides contextual information on Uruguayan worker-managed firms and describes the data; Section 2 presents the main results. Section 3 presents additional empirical results. Section 4 concludes and discusses the implications of these results in terms of the organisational performance of WMFs.

1. Context and Data

1.1. *Worker-managed Firms in Uruguay*

In Uruguay, WMFs are those firms that are legally registered as producer cooperatives (PCs) in which the employee-to-member ratio does not exceed 20%. Worker-managed firms are allowed to hire temporary employees in response to seasonal demand changes but they must still comply with the legislated maximum level of hired workers in order to receive certain tax advantages – in particular, the exemption from paying the employer payroll tax to social security.[4] The law also requires a minimum of six members to register a new cooperative firm.

Although their key organisational features are predetermined by law, WMFs have discretion over a broad range of associational rules. With respect to governance structure, WMFs must have a general workers' assembly that selects a council to supervise the daily operations (the council, in turn, usually selects the managers). Each member has only one vote, regardless of his capital contribution to the firm. Physical assets of WMFs can be owned by their members either collectively or individually. Under collective ownership, members do not own tradable shares but enjoy the right to usufruct as long as they work in the firm. Under individual ownership, members own capital shares that vary with the firm's value. Most Uruguayan WMFs operate under a collective ownership regime. As in other countries, membership markets are extremely rare in Uruguay. A recent survey indicates that less than 10% of Uruguayan WMFs are owned by their workforce through individual shares (Alves *et al.*, 2012).

[3] Abramitzky (2008) shows that more educated individuals and those employed in high-skilled occupations have a greater propensity to exit equal-sharing kibbutzim.

[4] There is no difference in the personal income tax regime applied to members of WMFs and employees of conventional firms.

1.2. *Worker-level Panel Data*

To test whether redistribution actually takes place within WMFs, I use a random sample of Uruguayan workers who were registered in social security at least one month during the period from January 1997 to April 2010. The data were provided by Banco de Prevision Social, the agency in charge of social security affairs in Uruguay. Employers are obliged to deliver monthly information on their employees to the agency, which uses that information to calculate pension and social benefits. The structure of the data is an unbalanced panel of workers extending from January 1997 to April 2010. The data contains information on daily wages, personal characteristics of the worker (gender, age, tenure), and attributes of the firm in which he works (firm size, industry). Each worker-month observation is tagged with a firm identification number so that job changes (and any other work history discontinuity) can be observed. Most importantly, the data identify the legal form of the firm for each worker's employment spell. Thus, workers employed by WMFs are identified as those working in a firm registered as a PC. I restrict the sample to workers employed by non-agricultural private firms; public and rural workers are excluded. Finally, I trim the data by excluding observations with daily wages corresponding to the top and bottom 1% of the wage distribution.

The descriptive statistics are presented in Appendix Table A1. The resulting sample includes, on average, about 40,000 workers in each month. Those employed in WMFs amount to only some 3% of all workers. Average wages are higher in worker-managed than in conventional firms. However, the composition of the two groups is different: workers employed by WMFs are older than those employed by CFs and, in the latter case, the average firm size is smaller. Proportionately fewer women are employed by WMFs than by CFs, although female participation in the former has increased over the period.

To give a preliminary picture of the extent of redistribution within WMFs, I compute two standard inequality measures for workers employed by WMFs *versus* CFs.[5] Figure 1 plots the evolution of the Gini (panel (*a*)) and Theil indexes (panel (*b*)) of both daily and hourly wages among workers employed in each type of firms. As expected, wage inequality is systematically lower in WMFs. For instance, the Gini index of daily wages is, on average, 9.3 percentage points lower for workers employed by worker-managed than by conventional firms. Wage inequality computed using hourly wages is also lower in WMFs than in CFs.[6] Figure 2 provides further information that characterises the wage distribution in WMFs and CFs. Worker-managed firms seem to reduce not only pay dispersion but also pay skewness, thus improving the median worker's compensation relative to the mean. Both the mean-to median wage ratio (panel (*a*)) and the coefficient of wage skewness (panel (*b*)) are systematically lower among workers employed by WMFs *versus* CFs.

[5] In each year, only workers between the ages of 20 and 55 are considered.

[6] It is also worth noting from Figure 1 that there seems to be some tendency towards convergence in wage dispersion between WMFs and CFs. This tendency is consistent with several institutional changes that the Uruguayan labour market has experienced since 2005, such a sharp increase minimum wage, the introduction of mandatory collective bargaining and higher unionisation rates.

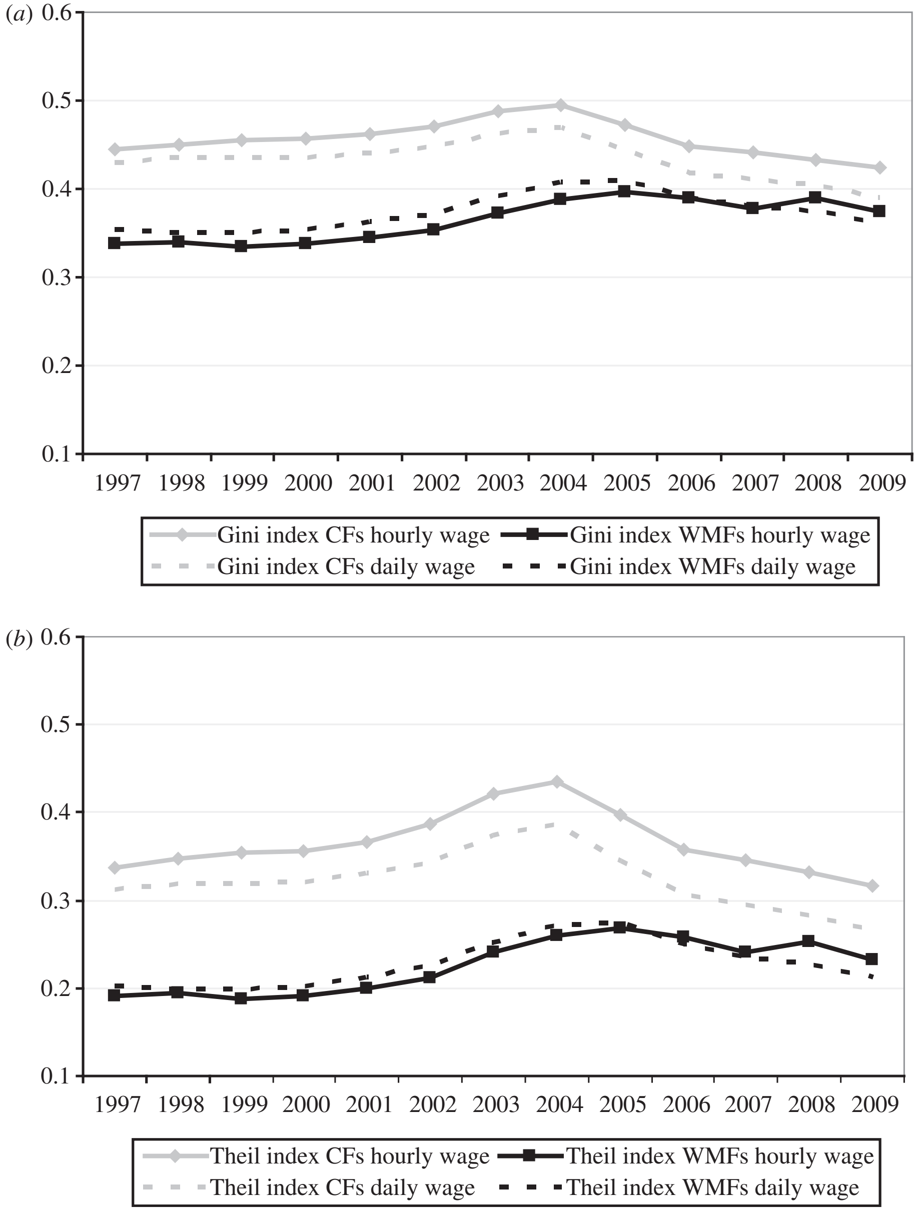

Fig. 1. *Wage Inequality in WMFs and CFs, 1997–2009*
Notes. Panel (*a*) reports the Gini index of both daily and hourly wages. Panel (*b*) reports the Theil index.

1.3. *Matched Organisation–Worker Panel Data*

To investigate whether WMFs suffer from brain drain and whether this problem is related with the extent of internal redistribution, I exploit a matched employer–employee monthly panel data set. The data cover the entire population of Uruguayan

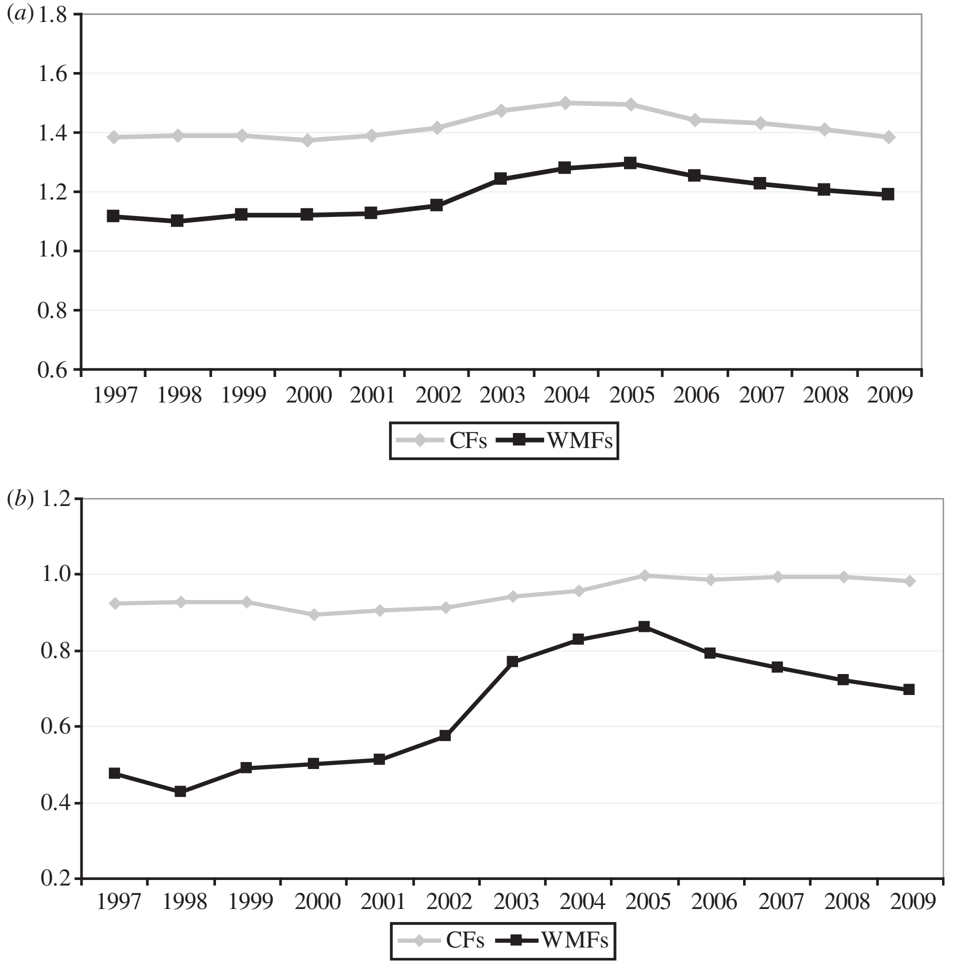

Fig. 2. *Mean-to-median Ratio and Wage Skewness in WMFs and CFs, 1997–2009*
Notes. Panel (*a*) reports the mean-to-median ratio of daily wages. Panel (*b*) reports the Pearson's coefficient of wage skewness, computed as [3 × (mean − median)] ÷ (standard deviation).

firms registered as producer cooperatives and all their workers (members and non-members) during the period from January 1997 to April 2010. Many PCs rely extensively on hired labour to carry out productive activities, which implies that – as in conventional firms – most of the workforce has no control over firm decisions. I therefore distinguish WMFs from the total population of PCs by using information of the employee-to-member ratio. I define WMFs as those PCs in which this ratio is lower than 20% at the time of entry. Estimates are performed using the subsample of WMFs just described.

The main advantage of the data is that it is possible, for each WMF, to match the information on all its workers in each month with a unique identification number. Hence the structure of the data is that of a linked employer–employee panel data set. Firm-level information includes firm size (measured as total employment) and industry

(5-digit SIC code). Worker-level information includes age, gender, job tenure, gross monthly wages, and number of days worked. Gross monthly wages are deflated by the consumer price index and divided by the number of days worked in order to obtain the real daily wage for each worker. I also exclude workers whose daily wages are outside the 1–99% range. Key to this study is that I can observe the entire wage distribution at any time and compute intra-firm pay dispersion indicators. The data enable me to observe each individual employment spell within WMFs and to locate workers' position in the firm's wage distribution. Among those workers who exit from WMFs during the period, I can also distinguish between voluntary quits and separation for other reasons (such as layoff, retirement, or death).

Descriptive statistics on workers and firms are reported in Appendix Tables A2 and A3, respectively. The resulting sample includes, on average, roughly 10,500 workers and 270 PCs in each month. Information on the subsample of WMFs is also presented. It is worth noting that average wages in the individual-based data (Table A2) are always higher than the average firm wage (Table A3). This difference simply reflects the fact that larger PCs, which account for more workers, have higher average wages than smaller PCs; that is, the (unweighted) average firm wage is disproportionately influenced by small, low-wage PCs.

2. Main Results

2.1. *Worker-managed Firms Redistribute in Favour of Low-wage Workers*

Section 1 gives *prima facie* evidence that inequality is lower among workers employed by WMFs than among those employed by CFs. Of course, that naïve comparison may be affected by the different workforce and sectoral composition of each firm type. To provide more systematic evidence on redistributive policies in WMFs, I use the worker-level panel described in subsection 1.2 and proceed as follows. First, in order to determine the sign and magnitude of the wage differential between workers employed in worker-managed and conventional firms, I estimate a standard Mincerian equation as follows:

$$\ln w_{ijt} = \boldsymbol{x}_{ijt}\boldsymbol{\alpha} + \boldsymbol{z}_{jt}\boldsymbol{\beta} + C_{ijt}\delta + \lambda_t + \eta_i + u_{ijt}, \tag{1}$$

where $\ln\, w$ denotes the logarithm of real daily wages of an individual i in firm j at time t, the $\boldsymbol{x}$ are observed characteristics (gender, age, and tenure as well as quadratics in age and tenure) of the individual worker, the $\boldsymbol{z}$ are observed features (size, industry) of the enterprise j by which the individual is employed, and C is a dummy indicator variable that is set to 1 when worker i is employed by a WMF (and set to 0 otherwise); the λ_t are year fixed effects.[7] Unobserved factors affecting wages are represented by the terms u and η, where the latter denotes unobserved factors that vary across individuals but are fixed for a given individual over time. The wage differential is captured by the coefficient δ.[8]

[7] One drawback to using social security data is the lack of information on workers' education level.

[8] Pencavel *et al.* (2006) adopt a similar empirical approach and find that, in Italy, being employed by a WMF is associated with a negative wage gap.

Table 1

Wage Gap Between Workers Employed in WMFs and CFs

	OLS	FE	FE	FE	FE
	(1)	(2)	(3)	(4)	(5)
Coop	0.055**	0.027*	0.028*	0.092**	0.091**
	(0.011)	(0.015)	(0.016)	(0.038)	(0.038)
Female	−0.230***				
	(0.005)				
Age	0.060***	0.210***	0.212***	0.212***	0.211***
	(0.001)	(0.002)	(0.002)	(0.002)	(0.002)
Age^2	−0.001***	−0.001***	−0.001***	−0.001***	−0.001***
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
Tenure	0.047***	0.032***	0.030***	0.030***	0.030***
	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)
Tenure^2	−0.001***	−0.001***	−0.001***	−0.001***	−0.001***
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
Firm size (in logs)	0.153***	0.122***	0.099***	0.100***	0.099***
	(0.001)	(0.001)	(0.002)	(0.002)	(0.002)
Year fixed effects	Yes	Yes	Yes	Yes	Yes
Industry fixed effects	Yes	Yes	Yes	Yes	Yes
Observations	5,264,811	5,264,811	3,533,031	3,445,097	3,445,097

Notes. The dependent variable is the log of daily wage. Column (1) reports pooled OLS estimates; columns (2)–(5) report panel data fixed-effect estimates. The estimates reported in columns (3)–(5) exclude workers employed in firms with fewer than six workers. Estimates in columns (4) and (5) compare employees in CFs with members in WMFs (i.e. non-members are excluded). All estimates include a set of 13 year dummies and six industry dummies. The estimates in column (5) also include sectoral-specific year fixed effects. Standard errors (reported in parentheses) are clustered at the individual level. *Significant at 10%; **significant at 5%; ***significant at 1%.

I estimate (1) via pooled ordinary least-squares (OLS) and fixed-effect (FE) regressions. The latter strategy is feasible because there is mobility of workers between WMFs and conventional firms. Under the assumption that selection into the WMF status is based on unobserved but time-invariant individual characteristics, fixed-effect regressions yield an unbiased estimate of the wage gap. The fraction of workers who switch between WMFs and CFs is roughly 4%.[9] It is well known (e.g. from the literature on unions) that FE estimates of a relatively persistent status – as when there are only a small number of switchers – are more susceptible to attenuation bias due to measurement errors (Freeman, 1984; Card, 1996). However, measurement errors are of less concern in this study because the estimates rely on administrative data that are extremely unlikely to reflect either misreporting or miscoding.

Estimates are reported in Table 1. Column (1) reports the results of the pooled OLS estimate, according to which a worker employed by a WMF earns 5.5% more than one employed by a CF; this difference is highly significant. However, an OLS estimate of (1) may be biased if C and η are correlated – that is, if unobservable factors affecting the choice between working for a WMF or a CF are correlated with the determinants of earnings. Column (2) reports the results from a fixed-effect regression that yields consistent estimates for δ under arbitrary correlation between C and η. The wage gap is

[9] Roughly, 30% of workers' transitions between WMFs and CFs correspond to WMF-to-CF switches.

still positive (2.7%) and significant at the 10% level.[10] I perform an additional FE estimate that excludes workers employed in micro-enterprises (i.e. firms employing fewer than six workers). The results, which are reported in column (3) of Table 1, remain unchanged. The estimates so far have compared all workers employed in WMFs (members and non-members) with those employed in CFs. Results are qualitatively similar when considering only WMF members. The wage gap is slightly higher (9%) and highly significant (see column (4) of Table 1). This is plausible given that WMF members' compensation includes distributed dividends. Finally, to account for heterogeneous time effects across sectors, column (5) reports estimates that include sectoral-specific year fixed effects. Results are robust also to this modification.[11]

Having documented a positive wage premium associated with being employed in a WMF, I then ask whether this wage gap varies across the wage distribution. If WMFs actually implement redistributive policies, then we should expect the magnitude of the wage differential to be greater at the bottom of the wage distribution. In other words, the gain experienced by a worker who moves from a conventional firm to a worker-managed firm should be greater for low-wage than for high-wage workers. To perform this analysis, I use quantile regression to estimate the wag gap associated with being employed in a WMF at each quantile $\theta \in [0, 1]$ of the distribution of the log of daily wages of worker i in firm j during month t:

$$Quant_\theta(w_{ijt}|\cdot) = Coop_{ijt}\gamma_\theta + \boldsymbol{X}_{it}\boldsymbol{\beta}_\theta + \boldsymbol{Z}_{jt}\boldsymbol{\delta}_\theta, \tag{2}$$

where $Quant_\theta(w_{ijt}|\ \cdot)$ refers to the conditional quantile of the log of daily wages, $\boldsymbol{X}_{it}$ captures personal characteristics (gender, age, age squared, tenure, tenure squared), and $\boldsymbol{Z}_{jt}$ stands for firm attributes (firm size, industry); $Coop_{ijt}$ is a dummy variable set equal to 1 only if individual i is employed by a WMF. Table 2 reports the results of Pooled quantile regression estimates for the 0.2, 0.4, 0.6, and 0.8 quantiles.[12] As expected, the wage premium associated with being employed in a WMF declines along the wage distribution and becomes negative at the top. The wage premium for the 0.2 quantile is 18% as compared with a wage penalty of 4% for the 0.8 quantile.[13] Compensation policies within Uruguayan WMFs seem to strongly favour workers at the bottom of the distribution.

Pooled QR estimates compare individuals with different unobserved ability. For such a reason, I implement the approach recently proposed by Canay (2011) to control for unobserved heterogeneity in a quantile regression setting. Table 2 also reports the results of the resulting fixed effect QR estimates. Consistent with pooled estimates, the wage premium associated with being employed in a WMF is significantly declining in wages, reinforcing the idea that WMFs actually redistribute in favour of low-wage

[10] The Hausman test leads to a strong rejection of the null hypothesis that random effects yield consistent estimates (p = 0.000).

[11] I replicate the estimates reported in column (4) when including both month and year fixed effects. Alternatively, I try adding a linear time trend. I also perform estimates using the log of hourly wages (instead of the daily wage) as the dependent variable. Results are robust to all these modifications. Estimates using daily wages are preferred because information on working hours is missing for nearly a fifth of the sample.

[12] In Appendix Table A5, I report the results of quantile regressions for each year separately, pooling monthly workers' records in each year. Interquantile differences appear to be quite stable over the period. Results remain unchanged when the log of hourly wages is used as the dependent variable.

[13] The null hypothesis of no interquantile differences is rejected in all cases.

Table 2

Wage Gap Across the Wage Distribution. Quantile Regression Estimates

	Pooled QR				Fixed effects QR			
	q20	q40	q60	q80	q20	q40	q60	q80
Coop	0.184***	0.111***	0.044***	−0.043***	0.041***	0.036***	0.026***	0.023***
	(0.002)	(0.002)	(0.002)	(0.002)	(0.001)	(0.001)	(0.001)	(0.001)
Female	−0.187***	−0.211***	−0.245***	−0.277***	0.014***	0.004***	−0.004***	−0.0146***
	(0.001)	(0.001)	(0.001)	(0.001)	(0.000)	(0.000)	(0.000)	(0.001)
Age	0.041***	0.050***	0.060***	0.078***	0.199***	0.208***	0.214***	0.218***
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
Age^2	−0.001***	−0.001***	−0.001***	−0.001***	−0.001***	−0.001***	−0.001***	−0.001***
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
Tenure	0.051***	0.049***	0.0475***	0.046***	0.038***	0.029***	0.0258***	0.0215***
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
Tenure^2	−0.001***	−0.001***	−0.001***	−0.001***	−0.001***	−0.001***	−0.001***	−0.001***
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
Firm size	0.162***	0.150***	0.145***	0.147***	0.120***	0.120***	0.121***	0.123***
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
Test of interquantile differences								
20th = 40th	[0.000]				[0.000]			
20th = 80th	[0.000]				[0.000]			
40th = 80th		[0.000]				[0.000]		
Observations	5,264,811	5,264,811	5,264,811	5,264,811	5,264,811	5,264,811	5,264,811	5,264,811

Notes. The dependent variable is the log of daily wages. The *Coop* dummy variable is set equal to 1 only for workers employed in a PC. Firm size is measured as the log of total employment in each firm. All estimates include six industry and 13 yearly dummies. Bootstrapped standard errors (reported in parentheses) are based on 200 replications. *Significant at 10%; **significant at 5%; ***significant at 1%.

workers. Interestingly, the wage gap seems to be partially driven by selection, particularly at the bottom of the wage distribution. The comparison between pooled QR and FE-QR estimates indicates that the sign of the pooled QR bias varies across the wage distribution, suggesting that the pattern of sorting into WMFs is heterogeneous. Results are partly in line with the prediction of a standard Roy model of selection (Roy, 1951; Borjas, 1987). As the compensation structure in WMFs is more compressed than in the capitalist sector and, hence, returns to ability are lower, one would expect that low-ability workers self-select into WMFs. This seems to be true at the top but not at the bottom of the wage distribution. Indeed, anecdotal and survey evidence suggests that WMFs rely on different recruitment channels than CFs (e.g. recommendations from incumbent members, trial periods) in order to ensure ideological commitment and screen out low-ability applicants, mitigating adverse selection effects (Benham and Keefer, 1991; Burdin, 2013). In the next subsection, I investigate in greater detail the selection in exit from WMFs.

2.2. *High-ability WMF Members are More Likely to Quit*

In this subsection I test whether redistributive policies implemented by WMFs affect workers' flows. Specifically, I analyse whether the hazard of voluntary separation is greater for high-ability workers. To perform this analysis, I use the linked organisation–worker panel described in subsection 1.3. Because the study focuses on voluntary quits, I restrict the sample in several ways. First, I exclude workers older than 55 because they are probably considering retirement. Second, I do not consider separations caused by firm closures. Third, separations due to other reasons (e.g. layoffs, death) are treated as censored.[14] Finally, I drop left-censored spells – that is, individuals who were already working in a given firm at the beginning of the study period (January 1997). The problem of right-censored observations is handled by using duration analysis techniques.

In order to identify high-ability workers, I divide the workforce of each firm (at any given moment in time) into two groups: those with wages above and those with wages below the firm's median wage. As an alternative threshold, I use the 80th percentile of the intra-firm wage distribution. The intuition behind this procedure is to use the within-firm wage variation to rank workers according to their ability types. Controlling for other observables characteristics of the worker and the firm, I assume that the position of the worker in the internal wage scale is a reasonable proxy for their position in the ability distribution. This approach requires one to assume that workers' payoffs are increasing in their own types (Bartolucci and Devicienti, 2012).

Figure 3 plots non-parametric estimates of the survival function and the hazard function for job separations while distinguishing between high and low-wage workers. These functions are calculated for both the whole sample of workers employed in PCs (Figure 3*a*, *b*) and the subsample of WMF members (Figure 3*c*, *d*). The hazard of job separation is systematically higher for high-wage workers in both cases. The log-rank test clearly rejects the null hypothesis that the survivor functions of the two types of

[14] Voluntary quits constitute 72% of total worker separations. As expected, the fraction of voluntary quits increases (to 82%) when the analysis is restricted to members.

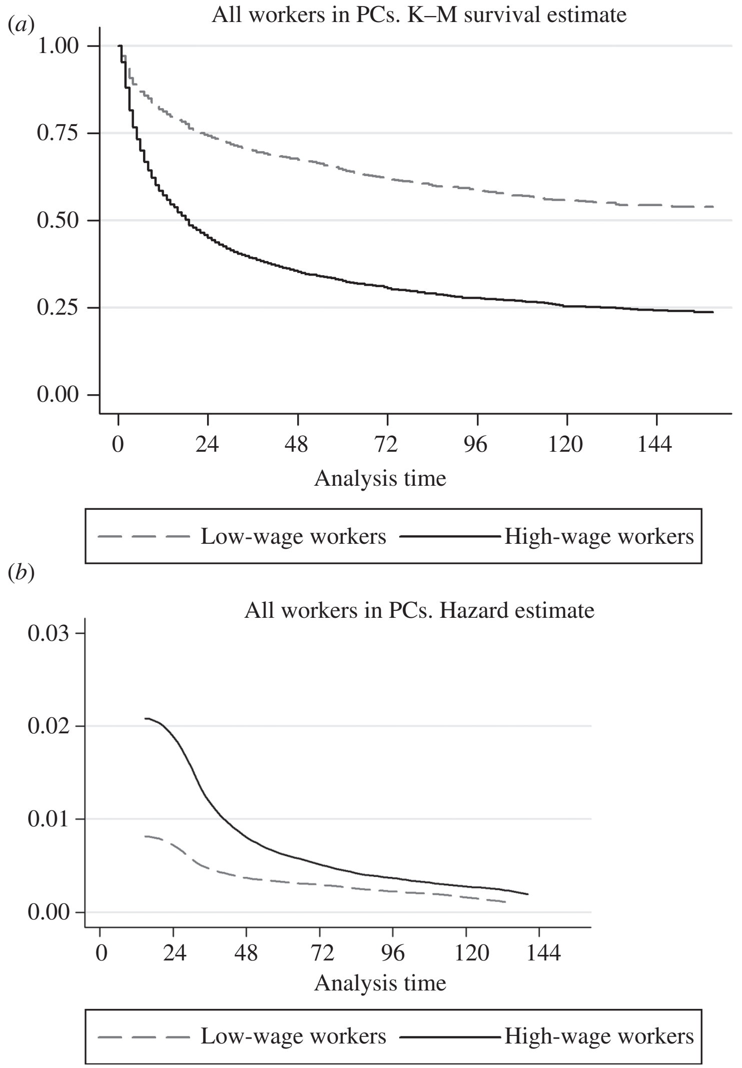

Fig. 3. *Worker's Position in the Within-firm Wage Distribution and Job Duration*
Notes. The High-wage worker indicator variable is set equal to 1 only for a worker whose daily wage is above the median daily wage in the firm that employs him. Figures 3*a*, *b* consider the full sample – that is, all workers (members and non-members) employed by PCs; in Figures 3*c*, *d* the estimates are restricted to members of WMFs. The Kaplan–Meier survivor function is defined as $\hat{S}(t_j) = \prod_{j|t_j<t}(1 - d_j/n_j)$, where d_j is the number of failures occurring at time t_j and where n_j is the number at risk at time t_j (before any failures occur). The hazard function is calculated as a weighted kernel density using the estimated hazard contributions: $\Delta\hat{H}(t_j) = \hat{H}(t_j) - \hat{H}(t_{j-1})$, where t_j is the current failure time and $\hat{H}(t_j)$ is the estimated cumulative hazard. The Nelson–Aalen estimator of $\hat{H}(t_j)$ is defined as $\hat{H}(t_j) = \sum_{j\,|\,t_j\leq t}(d_j/n_j)$; this is the sum of expected failures at each observed time. For further details on non-parametric survival analysis, see Jenkins (2005) and Cleves *et al.* (2008).

workers are equal ($\chi_{(1)}$ = 2,410). In order to analyse the determinants of employment duration in WMFs (i.e. the time elapsed between workers' enrolment and voluntary separation), I estimate a proportional hazard model (Cox, 1972):

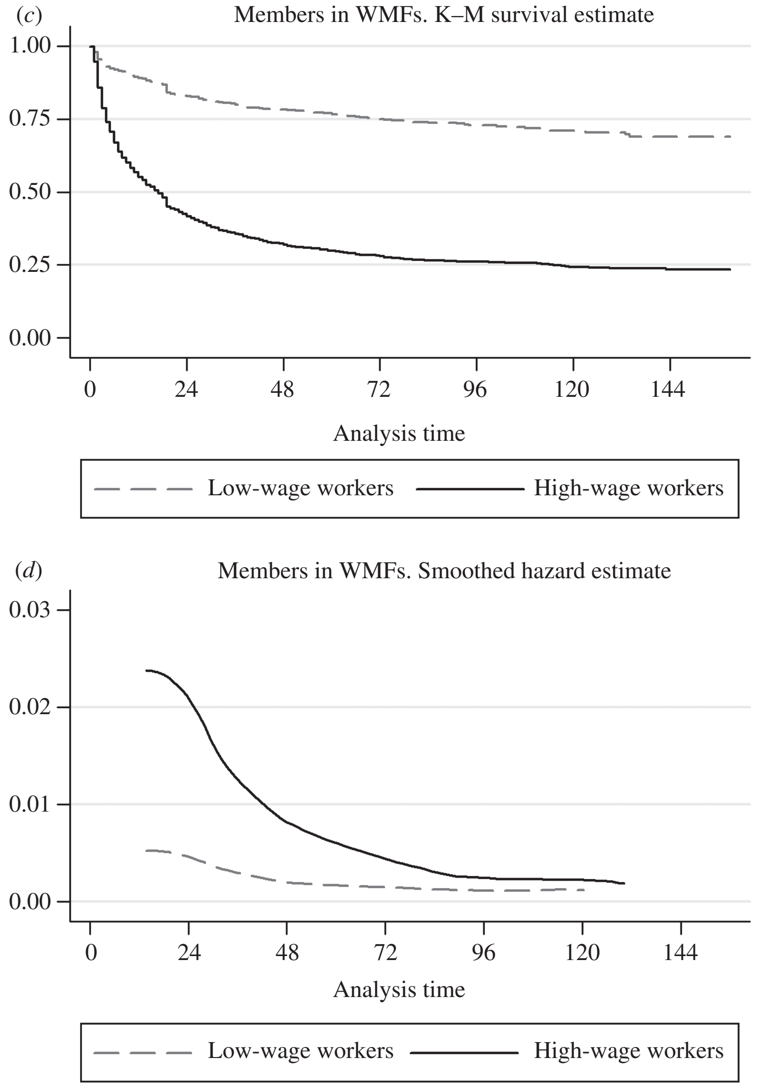

Fig. 3. *(Continued)*

$$h_{ij}(t) = h_j(t)\exp(HighW_{it}\beta_1 + \boldsymbol{X}_{it}\boldsymbol{\beta}_2 + \boldsymbol{Z}_{jt}\boldsymbol{\beta}_3), \tag{3}$$

where $h_j(t)$ is the baseline hazard for firm j and where t is the number of months that individual i has been employed at firm j; the dummy variable $HighW_{it}$ is set equal to 1 for workers whose daily wage is above the firm's median daily wage, $\boldsymbol{X}$ is a vector of personal characteristics (gender, age, age squared), and $\boldsymbol{Z}$ is a vector of firm characteristics (firm size, proportion of female workers, workforce average age and its dispersion). The effect of a unit change in any covariate is to produce a constant proportional change in the hazard rate. The coefficient of interest is β_1.[15] To rule out potential unobserved firm-level confounding factors, I estimate stratified Cox models in which each firm has its own flexible baseline hazard function. This approach allows one to control for all time-invariant firm-level characteristics (Giuliano *et al.*, 2011). Cox model estimates stratified by firm eliminate unobserved heterogeneity across firms but not across individuals within a firm. I account for unobserved individual-level heterogeneity by also estimating a parametric model in which each individual's duration depends on a random effect ('frailty') and the baseline hazard is assumed to have a log-normal distribution (Jenkins, 2005).[16]

Table 3 reports the results. All estimates are restricted to the subsample of members of WMFs. Column 1 reports the results of estimating (3) while controlling only for personal characteristics. In column (2) the estimates control also for firm-level characteristics and include cohort fixed effects to account for common shocks (at the time of entry) that may affect subsequent job duration. Column (4) reports estimates of the parametric frailty model. Results are qualitatively similar across specifications. The hazard of job separation is systematically greater for high-ability workers. The results reported in column (2) indicate that high-wage members are 3.7 times more likely than are low-wage members to exit.[17] In estimates reported in column (3), the variable *HighW* equals 1 if the individual's daily wage is above the 80th percentile of their intra-firm wage distribution. Under this alternative definition, the hazard of exit of high-ability workers is 4.5 times higher than the hazard of low-ability ones. Estimates reported in column (4), which account for individual unobserved heterogeneity, indicate that the time ratio associated with being a high-ability worker is 0.23; this means that the status of high-ability member reduces employment duration (survival time) within a WMF by 77%, or roughly 20 months.[18] That the high-ability individuals are more likely to exit provides further support for the idea that pay compression in WMFs is a deliberate policy. As Lazear and Shaw (2009) point out, there would be no reason for top workers to leave disproportionately (nor for bottom workers to stay disproportionately) if all workers were paid their competitive wage.[19]

[15] The Breslow method is used for handling ties. I check the proportional hazard (PH) assumption by means of graphical methods (Jenkins, 2005). This assumption seems to be satisfied by the data; see Appendix Figure A1. I also perform the test based on the Schoenfeld residuals for the variable *HighW* and do not reject the PH assumption ($p = 0.218$). The PH assumption is not rejected (at 5%) when the global test of the model is considered ($p = 0.077$).

[16] The log-normal distribution is consistent with the non-monotonic pattern of duration dependence of the hazard observed in Figure 3. Unlike the Cox model, the log-normal model does not rely on the PH assumption.

[17] By expressing the model in terms of the log of the hazard ratios, this effect is computed as $\exp(1.32)$.

[18] This effect is computed as $[1 - \exp(-1.484)] \times 100 = 77.32$. The mean employment duration for the subsample of WMF members is 27 months; thus, $(27 \times 0.77)/12 = 1.73$. Observe that, in column (3), the covariate effects must be interpreted in terms of survival time ('accelerated failure time' metric) and not in terms of the hazard as in Cox model estimates ('proportional hazard' metric).

[19] I perform additional robustness checks as well. First, I estimate the Cox model considering all workers (members and non-members) in WMFs. Second, I consider the whole sample of workers employed in all PCs. Third, I exclude employment spells with time gaps. Finally, I estimate the Cox model defining the worker's position in the within-firm wage distribution at the time of entry. None of the described modifications alters the basic results. Estimates for these alternative regressions are available from the author upon request.

Table 3

Worker's Position in the Within-firm Wage Distribution and Hazard of Exit in WMFs. Results from Duration Models Estimates

	(1)	(2)	(3)	(4)	(5)	(6)	(7)
(A) *HighW*	1.320***	1.307***	1.512***	−1.484***	1.895***	2.398***	1.375***
	(0.0529)	(0.0529)	(0.0466)	(0.0614)	(0.123)	(0.190)	(0.0690)
(B) *HighW* × *Coef. of variation*					−1.610***		
					(0.254)		
Coef. of variation					1.606***		
					(0.258)		
(C) *HighW* × *Mean-to-median ratio*						−0.995***	
						(0.149)	
Mean-to-median ratio						1.855***	
						(0.164)	
(D) *HighW* × *Founding member*							−0.428***
							(0.119)
Founding member							−0.251**
							(0.119)
Hazard ratio/Time ratio							
(A)	3.743***	3.695***	4.536***	0.227***			3.955***
	(0.198)	(0.196)	(0.212)	(0.014)			(0.273)
Post-estimation: (A) + σ*(B)					2.482***		
					(0.177)		
Post-estimation: (A) − σ*(B)					5.054***		
					(0.434)		
Post-estimation: (A) + σ*(C)						3.274***	
						(0.177)	
Post-estimation: (A) − σ*(C)						4.129***	
						(0.268)	
Post-estimation: (A) + (D)							2.579***
							(0.254)
Worker-level controls	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Firm-level controls	No	Yes	Yes	Yes	Yes	Yes	Yes
Cohort fixed effects	No	Yes	Yes	Yes	Yes	Yes	Yes
Observations	183,523	183,514	183,514	183,514	163,151	163,151	96,722

Notes. Cox proportional hazard models stratified by firm – except for column (4), which reports estimates from a shared 'frailty' model in which the baseline hazard assumes a log-normal distribution. The *HighW* dummy variable is set equal to 1 for those workers whose daily wage is above the firm's median daily wage (and to 0 for other workers), except in estimates reported in column (3) in which *HighW* is set equal to 1 for workers whose daily wage is above the 80th percentile of their intra-firm wage distribution. All estimates control for worker-level characteristics (gender, age, age squared) and are restricted to WMF members. Estimates presented in columns (2)–(6) also control for firm-level characteristics (firm size, average age of the workforce and its dispersion, fraction of female workers) and cohort fixed effects. The estimates presented in column 4 include industry fixed effects; in columns (4) and (5), the estimates are restricted to WMFs employing at least 10 workers. In column (7), estimates are restricted to WMFs (formed after January 1997) for which founding members can be identified. Robust standard errors (reported in parentheses) are adjusted for clustering at the individual level. *Significant at 10%; **significant at 5%; ***significant at 1%.

2.3. *High-ability Workers are Less Likely to Quit when Redistribution Within WMFs is Less Intense; Founding Members are Also Less Prone to Exit*

One major concern about previous results is that greater mobility might be a general feature of high-ability individuals and not a specific feature of WMFs. To assess whether the degree of equality within WMFs affects the outflow of high-ability members, I am able to exploit the observed within-firm variation in intra-firm wage dispersion among WMFs. As already mentioned, this procedure allows me to estimate models that control for unobserved differences across firms. The expectation is that a less compressed wage structure mitigates brain drain. To test this hypothesis, I estimate (3) while including a measure of intra-firm inequality and its interaction with the variable identifying high-wage members. Because measures of intra-firm inequality are not meaningful for small firms, I restrict the sample to WMFs employing at least ten workers. In order to characterise the wage distribution within each WMF, I consider two measures: the coefficient of variation and the mean-to-median ratio of wages within the firm.[20] I expect the coefficient for the interaction term to be negative. If brain drain is driven by egalitarian wage policies implemented by WMFs then, *ceteris paribus*, high-wage workers should be less likely to exit WMFs in which redistribution is less pronounced.

The results reported in columns (5) and (6) of Table 3 support this hypothesis. The interaction term is negative and statistically significant in both specifications, and the magnitude of the effect is sizable. I report the post-estimation of the hazard ratio (using a linear combination of parameter estimates) when the within-firm coefficient of variation in daily wages is one standard deviation (σ) above or below the mean. According to the values in column (5) of the Table, the hazard ratio of high-ability members is twice as high in a WMF for which the within-firm coefficient of variation in wages is one standard deviation (0.221) below the sample mean (0.392). Results are qualitatively similar in estimates that include the mean-to-median firm wage ratio (see column (6) of Table 3). The hazard ratio of high-ability members is 1.26 times higher in a WMF for which the mean-to-median wage ratio is one standard deviation (0.117) below the sample mean (1.101). It should be emphasised that the mean-to-median wage ratio has a direct interpretation in terms of a WMF median voter model (Kremer, 1997). Higher values of the mean-to-median ratio indicate that the median voter commits not to engage in redistribution while taking into account participation constraints of the most productive members. A consistent feature of the findings reported here is that the brain drain effect is mitigated in those WMFs whose median member is less prone to leverage his pivotal position in the organisational political process to redistribute away from high-ability members.

Finally, I analyse whether the hazard ratio of high-ability members varies with their status in the organisation. Previous evidence from Israeli kibbutzim indicates a positive association between the degree of equality and the degree of members' ideology (Abramitzky, 2008). Ideology seems to play the role of relaxing the participation constraint by increasing the non-pecuniary value of staying in the kibbutz. It is

[20] I compute the average of these variables over each individual employment spell. However, the averages vary both between and within firms, they vary only between (not within) individuals. In this way I can estimate the Cox model stratified by the firm.

Table 4

Labour Market Conditions and Hazard of Exit in WMFs. Results from Duration Models Estimates

	(1)	(2)	(3)	(4)
HighW	1.468***	1.531***	1.530***	1.709***
	(0.061)	(0.070)	(0.070)	(0.067)
Ratiow		0.207**	0.210***	0.095
		(0.083)	(0.082)	(0.066)
Unemp	−0.039***		−0.039***	−0.012
	(0.014)		(0.014)	(0.014)
HighW × *Unemp*				−0.089***
				(0.011)
HighW × *Ratiow*				0.256***
				(0.094)
Worker-level controls	Yes	Yes	Yes	Yes
Firm-level controls	Yes	Yes	Yes	Yes
Observations	163,949	159,628	159,628	158,917

Notes. Cox proportional hazard models stratified by firm. The *HighW* dummy variable is set equal to 1 for those workers whose daily wage is above the firm's median daily wage (and to 0 for other workers); *Ratiow* is the ratio of the median daily wage corresponding 2-digit sector of the WMF to the member's daily wage; and *Unemp* is the monthly urban unemployment rate. All estimates include *Ratiow* and *Unemp* (lagged three months) and are restricted to WMF members. In addition, all estimates control for worker-level characteristics (gender, age, age squared) and firm-level characteristics (firm size, average age of the workforce and its dispersion, fraction of female workers). Robust standard errors (reported in parentheses) are adjusted for clustering at the individual level. *Significant at 10%; **significant at 5%; ***significant at 1%.

unfortunate that I have no direct measure of a member's ideology. Nonetheless, it is possible to identify the founding members of WMFs formed after January 1997. It is reasonable to assume that the ideological commitment of first-generation members is stronger than that in subsequent members. I estimate (3) while including an indicator variable for the founding member and its interaction with the variable identifying high-wage members. These results are reported in column (7) of Table 3. On average, founding members are less likely to quit WMFs. A finding of particular interest is that the hazard ratio of high-ability members is 1.4 times lower in the case of founding members. This result confirms the intuitive notion that redistributive policies of WMF are less constrained by the threat of brain drain when members are intrinsically motivated to join the firm.[21]

2.4. *High-ability Members are Less Likely to Quit When Outside Options are Less Attractive*

Finally, I analyse whether the hazard of exit of high-ability members varies according to changes in labour market conditions in the capitalist sector. To characterise the external labour market, I use three-month lagged values of both the monthly urban unemployment rate, $Unemp_{t\text{-}3}$, and the ratio of the median daily wage paid in the capitalist sector – computed for the specific 2-digit sector of the WMF in which the

[21] First-generation members may also have greater sunk investments in their firms. Therefore, I cannot rule out that a founding member's lower hazard of exit is due to lock-in effects associated with the collective ownership of a WMF's physical assets. Indeed, Abramitzky (2008) finds that the degree of equality is higher in wealthy kibbutzim and that higher wealth reduces the brain drain in equal-sharing kibbutzim.

individual is employed – to the member's daily wage, $Ratiow_{it-3}$. I then estimate (3) while including these variables and their interaction with the variable identifying high-wage members within WMFs.

Results are reported in Table 4. As expected, the more (less) attractive are the external labour market conditions, the higher (lower) is the hazard of exit in WMFs. Column (4) of the Table reports estimates that include labour market conditions interacted with the variable $HighW_{it}$, which identifies high-wage members within WMFs. Both interaction terms have the expected sign and are highly significant. It is worth noting that the sensitivity of quit decisions to external labour market conditions also varies according to the member's position in the intra-firm wage distribution. When outside options in the capitalist sector become more attractive, the exit hazard increases more for high-ability than for low-ability members.

3. Additional Empirical Results

3.1. *Mobility of High-ability Workers Employed in the Conventional Sector*

In subsection 2.2, I simply compare the hazard rate between high and low-wage members in WMFs. Unfortunately, the matched organisation-individual data set described in subsection 1.3 only provides information of WMFs and their workers, and, hence, precludes studying the quit behaviour of workers employed at conventional firms during the same period. The lack of a suitable control group is an obvious limitation of that approach. One can certainly argue that high-ability workers are more mobile in any organisational setting and so not simply because of redistributive policies implemented by WMFs.

To address this major concern, I conduct two further empirical analyses. First, I analyse how CF-to-CF transitions correlate with workers' pre-exit wage using the panel of workers described in subsection 1.2. In Appendix Table A6, I report estimates from binary outcome models in which the dependent variable is a dummy that takes a value of one if the worker switched from a CF to another CF. Results from both Probit and Logit regressions indicate that the probability of switching from a CF to another CF is negatively correlated with the worker's pre-exit wage. Results are qualitatively similar when instead of the log of pre-exit daily wage I use wage quintiles to indicate the position of the worker in his sector-specific wage distribution. The probability of being a CF-to-CF switcher is lower for workers located at the top of their sector-specific wage distribution.

Second, I conduct a similar analysis but using a matched employer–employee panel containing annual information of the universe of Uruguayan formal workers and their firms over the period 2009–12. The data comes from administrative tax records provided by the National Tax Agency in Uruguay (*Dirección General Impositiva*). Even though available for a shorter and more recent time period, this data set allows to implement exactly the same empirical strategy adopted in subsection 2.2 but for workers employed in the capitalist sector. This consists in ranking both movers and stayers in terms of ability according to their position in their intra-firm wage distribution. High-ability workers are defined as those receiving a wage above the 80th percentile of their firm's wage distribution.

I restrict the sample in several ways. First, I exclude workers whose wages are outside the 1–99% range. Second, I focus on transitions between main jobs, defined as the job that is the primary source of earnings in the year. Third, I exclude workers under 20 and over 55 years old and those who were employed in the public sector at some point during the period. Fourth, I do not consider individuals who separate from a distressed firm, that is, a firm that experienced a 40% or larger employment loss in the year. The purpose of all these exclusions is to focus on voluntary job-to-job transitions and workers with a strong labour market attachment. Descriptive statistics summarising the resulting sample are reported in Appendix Table A7.

I use these data to investigate the relationship between workers' separation decisions and their position in their intra-firm wage distribution. Results are reported in Table 5. Columns (1)–(3) report results from probit estimates in which the dependent variable is equal to 1 if the individual separates from their firm during the period and 0 otherwise. To test whether results are robust to alternative distributional assumptions, column (4) presents results from logit estimates. Finally, column (5) reports estimates from a random effect complementary log-log model which assumes the unobserved heterogeneity term (frailty) to be normally distributed. In columns (3)–(5), estimates are restricted to individuals without gaps in their employment history during the period. This is done in order to exclude job-to-unemployment transitions. In contrast to previous findings on WMFs, high-wage individuals in CFs are less likely to quit their firms than low-wage ones. This result is robust across all specifications.

3.2. *Post-exit Wage*

In this subsection, I investigate whether those individuals who leave a WMFs actually do better after quitting. I estimate a Mincer-type wage regression similar to (1), using the matched employer–employee 2009–12 panel described in the previous subsection, but restricting the analysis to those individuals initially employed at a WMF:[22]

$$\ln w_{ijt} = \boldsymbol{x}_{ijt}\boldsymbol{\alpha} + \boldsymbol{z}_{jt}\boldsymbol{\beta} + Mover_{ijt}\gamma_1 + Mover_{ijt} \times HighW_i^{09}\gamma_2 + \lambda_t + \eta_i + u_{ijt}, \qquad (4)$$

where $\ln w$ denotes the logarithm of real annual wages, $\boldsymbol{x}$ are observed characteristics of the individual worker (gender, age, and tenure as well as quadratics in age and tenure), $\boldsymbol{z}$ are observed features of the enterprise j by which the individual is employed (firm size, average age of the workforce and its dispersion, and fraction of female workers), *HighW* is a dummy that takes value of 1 if the worker's compensation exceeds the 50th percentile of their firm's wage distribution in 2009 (or, alternatively, the 80th percentile) and *Mover* is a dummy variable equal to 1 if the worker switched firms between $t-1$ and t; λ_t and η_i are year and individual fixed effects. The wage change experiences by movers relative to stayers is captured by the coefficient γ_1, and the differential effect for high-wage individuals is given by the coefficient γ_2 attached to the interaction term. The implied total wage change for high-wage individuals is $\gamma_1 + \gamma_2$.

[22] The matched organisation-individual data set used to perform duration models estimates (reported in subsection 2.2) does not allow to track individuals after quitting a WMF.

Table 5

Worker's Position in the Within-firm Wage Distribution and Probability of Exit from CFs

	Probit	Probit	Probit	Logit	Random effect cloglog
	(1)	(2)	(3)	(4)	(5)
HighW	−0.183***	−0.157***	−0.139***	−0.144***	−0.083***
	(0.002)	(0.002)	(0.002)	(0.002)	(0.001)
Female	−0.070***	−0.058***	−0.057***	−0.058***	−0.022***
	(0.002)	(0.002)	(0.002)	(0.002)	(0.001)
Age	−0.023***	−0.030***	−0.034***	−0.032***	−0.008***
	(0.001)	(0.001)	(0.001)	(0.001)	(0.000)
Age^2	0.000***	0.000***	0.000***	0.000***	0.000***
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
% Female	0.028***	0.016***	0.033***	0.035***	0.022***
	(0.004)	(0.004)	(0.004)	(0.004)	(0.002)
Average workforce age	−0.010***	−0.009***	−0.009***	−0.009***	−0.003***
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
Dispersion of workforce age (SD)	0.011***	0.010***	0.009***	0.009***	0.002***
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
Firm size	−0.013***	−0.013***	−0.016***	−0.016***	−0.007***
	(0.000)	(0.000)	(0.001)	(0.001)	(0.000)
Lnt					−1.629***
					(0.013)
ln σ_v					−11.262***
					(1.743)
ρ					7,82e-06
					(0.000)
Observations	324,749	324,749	227,208	227,208	613,419

Notes. Estimates based on the matched employer–employee 2009–12 panel. Columns (1)–(4) report average marginal effects from binary outcome models estimates in which the dependent variable is equal to 1 if the individual separates from their firm during the period and 0 otherwise. Column (5) reports estimates from a random effect complementary log–log model which assumes the unobserved heterogeneity term (frailty) to be Normally distributed. Estimates presented in columns (1) and (2) are performed using the whole sample while those reported in columns (3) and (4) are performed using the balanced panel of individuals. Estimates are restricted to individuals aged 20–55. Separations from distressed firms are not considered. The *HighW* dummy variable is set equal to 1 for those workers whose annual wage is above the 80th percentile of their intra-firm wage distribution (and to 0 for other workers), except in column (1) in which is equal to 1 for those workers whose wage is above the 50th percentile; all estimates control for worker-level characteristics (gender, age, age squared) and firm-level characteristics (firm size, average age of the workforce and its dispersion, fraction of female workers, industry dummies). Robust standard errors reported in parentheses. *Significant at 10%; **significant at 5%; ***significant at 1%.

Estimates are restricted to individuals employed at a WMF in 2009. Switchers from distressed firms are not considered.

Estimates of (4) are reported in Table 6. Columns (1) and (3) report results from Pooled OLS estimates. Columns (2), (4), and (5) report results from spell fixed effects estimates in which potential confounding factors associated with both individual and firm-level unobserved heterogeneity are jointly removed. On average, movers experience a wage loss compared to stayers in WMFs. However, the wage effect of switching firms appears to be significantly positive for individuals whose wages exceed the 80th percentile of their WMF wage distribution. Results reported in columns (3)–(5) indicate that high-wage workers who leave a WMF actually experience a gain of approximately 7–9% in their wages upon quitting.

Table 6

(Log) Wage Regression. Movers Versus Stayers in WMFs

	High-wage workers w > p50		High-wage workers w > p80		
	Pooled OLS	FE	Pooled OLS	FE	FE
	(1)	(2)	(3)	(4)	(5)
Mover	−0.816***	−0.225***	−0.792***	−0.310***	−0.337***
	(0.083)	(0.071)	(0.071)	(0.066)	(0.065)
Mover × *HighW*	0.319**	−0.161	0.865***	0.401**	0.409**
	(0.135)	(0.134)	(0.190)	(0.156)	(0.167)
Female	−0.244***		−0.244***		
	(0.032)		(0.032)		
Age	0.089***	−0.422***	0.089***	−0.424***	−0.460***
	(0.010)	(0.016)	(0.010)	(0.016)	(0.017)
Age^2	−0.001***	−0.001***	−0.001***	−0.001***	−0.001***
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
% Female	−0.289***	0.372*	−0.285***	0.364*	0.315
	(0.075)	(0.194)	(0.075)	(0.190)	(0.204)
Average workforce age	−0.007*	−0.002	−0.007*	−0.002	0.001
	(0.004)	(0.004)	(0.004)	(0.004)	(0.005)
Dispersion of workforce age (SD)	−0.048***	−0.001	−0.047***	−0.001	0.001
	(0.007)	(0.006)	(0.007)	(0.006)	(0.006)
Firm size	0.278***	0.309***	0.279***	0.323***	0.333***
	(0.007)	(0.057)	(0.007)	(0.055)	(0.058)
Industry fixed effects	Yes	Yes	Yes	Yes	Yes
Year fixed effects	Yes	Yes	Yes	Yes	Yes
Worker-firm fixed effects	No	Yes	No	Yes	Yes
Observations	13,451	13,451	13,451	13,451	11,103

Notes. Estimates based on the matched employer–employee 2009–12 panel. Estimates are restricted to individuals aged 20–55 who were employed at WMFs in 2009. Separations from distressed firms are not considered. In columns (1)–(2), the *HighW* dummy variable is set equal to 1 for those workers whose annual wage is above the 50th percentile of their intra-firm wage distribution (and to 0 for other workers). In columns (3)–(5), the *HighW* dummy variable is set equal to 1 for those workers whose annual wage is above the 80th percentile of their intra-firm wage distribution (and to 0 for other workers). All estimates control for worker-level characteristics (gender, age, age squared), firm-level characteristics (firm size, average age of the workforce and its dispersion, fraction of female workers), industry and time fixed effects. Estimates reported in columns (2), (4) and (5) control for worker-firm fixed effects. Estimates reported in column (5) are performed using the balanced panel of individuals. Robust standard errors (reported in parentheses) are adjusted for clustering at the individual level. *Significant at 10%; **significant at 5%; ***significant at 1%.

3.3. *Selection and Redistribution*

Workers may be self-selected into WMFs according to unobservable traits and preferences (e.g. risk and inequality aversion) that might also affect intra-firm pay dispersion (Ben-Ner and Ellman, 2013). Despite the extensive use of panel data techniques made throughout this article, it is impossible to control for all the potential confounding factors involved in the sorting process of individuals into WMFs in a non-experimental setting. In any case, the formation of a WMF is always mediated by individual choices. Hence, one should be cautious in interpreting the negative correlation between workplace democracy and pay dispersion reported in this article as a causal relationship.

However, additional survey and experimental evidence seems to support the idea that workers' control reduces pay dispersion. First, available survey evidence comparing Uruguayan WMFs created from scratch with those created through transformation of bankrupted conventional firms (*empresas recuperadas*) shows no significant differences in compensation structure and supervision intensity between the two creation modes (Burdin, 2013). Both types of WMFs exhibit lower pay dispersion and supervision intensity than comparable conventional firms. This is interesting because one could argue that non-random selection is more likely to occur in WMFs created from scratch as members deliberately choose to start-up a particular type of organisation. While it is not possible to rule out selection effects completely in the case of WMFs created from the transformation of a pre-existing conventional firm, it is plausible to expect that the formation of a WMF in this situation is more likely to be driven by an exogenous shock (firm bankruptcy) than by workers' preferences. Second, experimental studies in which selection plays no role – because random assignment guarantees that the allocation of subjects between treatment and control is fully exogenous – show that democratic institutions affect cooperative behaviour (Dal Bó *et al.*, 2010), organisational decision processes affect ethical behaviour towards outsiders (Ellman and Pezanis-Christou, 2010) and group identity significantly affects preference over redistribution (Klor and Shayo, 2010). The few attempts to study workplace democracy using economic experiments also report positive effects on effort, perceived pay fairness and individuals' satisfaction with a floor-constraint principle of distributional justice (Frohlich and Oppenheimer, 1990; Frohlich *et al.*, 1998; Mellizo *et al.*, 2014).

4. Conclusion and Discussion

In this article, I study the extent and effects of redistribution in WMFs. The analysis supports two main findings. First, WMFs exhibit a more compressed compensation structure than conventional firms. There is only a small wage premium associated with being employed in a WMF, and this gap declines significantly with increasing wage. This result is consistent with the hypothesis that workplace democracy is associated with substantial redistribution among workers. Second, WMFs suffer from brain drain: the separation hazard of high-ability members is more than three times higher than that of low-ability members. Moreover, I find that there is a relationship between the extent of pay compression and the severity of brain drain in WMFs: high-ability workers are less likely to exit a WMF whose wage structure is less compressed. In addition, the status of founding member is generally associated with a lower hazard of exit and significantly reduces the hazard of high-ability members, suggesting that the presence of intrinsically motivated workers enables greater redistribution. Finally, I find that the quit behaviour of high-ability members varies as a function of labour market conditions in the capitalist sector. By using a shorter matched employer–employee panel, I show that individuals located at the top of the intra-firm wage distribution in CFs are less likely to quit and that high-wage workers (employed at a WMF in 2009), who switch to the capitalist sector, experience a gain in their wages upon quitting. This reinforces – at least in the very specific Uruguayan context – the idea that an interplay between pay compression and quit behaviour of high-ability individuals in WMFs exists.

It is beyond the scope of this article to analyse the relationship between pay compression and organisational performance in WMFs. The brain drain effects documented here suggest a plausible mechanism to account for a potential negative relationship between pay compression and performance. Another possible explanation, which is suggested by tournament theory, is that a compressed wage structure reduces the expected gains from internal promotions and hence does not provide enough incentive to increase workers' efforts (Lazear and Rosen, 1981). Nevertheless, panel data evidence on the relative efficiency of WMFs indicates that they perform as well as (or even better than) conventional firms in terms of productivity (Craig and Pencavel, 1995; Fakhfakh *et al.*, 2012; Pencavel, 2013). Burdin (2014) also shows that Uruguayan WMFs are less likely to dissolve than are conventional firms. Experiments on team production also find positive performance effects associated with workplace democracy (Frohlich *et al.*, 1998; Mellizo *et al.*, 2014). Those experimental and non-experimental studies suggest that other beneficial effects associated with pay compression are at work in WMFs. The costs of equality associated with brain drain and inferior management quality may be outweighed by other labour discipline benefits, such as higher motivation of shop-floor workers, greater workplace cooperation and lower supervision costs (Milgrom and Roberts, 1990; Levine, 1991; Encinosa *et al.*, 2007).[23] Further research is needed to investigate the potential efficiency-enhancing effects of pay compression in democratically controlled workplaces.

[23] Survey evidence indicates that supervision intensity is significantly lower in Uruguayan WMFs than in conventional firms. WMFs rely more on mutual monitoring among co-workers to ensure workplace discipline. Interestingly, egalitarian WMFs have lower supervision intensity and rely more on horizontal monitoring than non-egalitarian WMFs (Burdin, 2013).

Appendix A. Descriptive Statistics and Additional Estimates

Table A1

Descriptive Statistics. Panel of Workers

	1997		2001		2005		2009	
	CF	WMF	CF	WMF	CF	WMF	CF	WMF
Number of workers	36,117	1,305	33,944	1,092	38,148	1,138	46,667	1,220
Fraction female workers	0.43	0.36	0.45	0.38	0.45	0.42	0.45	0.44
Age	36.34	41.21	37.32	42.66	37.59	43.51	38.07	43.09
	(12.63)	(10.57)	(12.27)	(10.75)	(12.16)	(10.80)	(12.18)	(11.13)
Tenure	5.26	9.12	5.80	10.33	5.39	10.81	4.87	10.15
	(6.67)	(8.12)	(6.82)	(8.64)	(6.97)	(9.28)	(6.62)	(9.76)
Monthly wage	13,829	25,138	13,118	22,632	10,779	17,880	13,376	21,210
	(13,260)	(17,546)	(13,398)	(15,693)	(11,181)	(15,551)	(12,428)	(16,028)
Daily wage	523.55	922.90	497.15	911.43	416.68	668.41	519.10	804.54
	(469.08)	(592.96)	(469.11)	(620.86)	(394.11)	(520.80)	(434.06)	(543.61)
Hourly wage	89.60	156.81	87.65	143.53	71.99	115.60	89.60	131.68
	(84.72)	(103.06)	(86.34)	(90.87)	(75.82)	(92.15)	(85.70)	(95.99)
Firm size	3.74	5.78	3.81	5.68	3.81	5.45	3.94	5.69
	(1.96)	(1.77)	(2.02)	(1.76)	(2.01)	(1.74)	(2.03)	(1.76)
Fraction in manufacturing	0.29	0.29	0.23	0.26	0.24	0.29	0.22	0.22
Fraction in transport	0.07	0.30	0.08	0.31	0.08	0.27	0.08	0.25
Fraction in services	0.32	0.40	0.36	0.41	0.36	0.41	0.35	0.49

Notes. Summary statistics reported in October of each year. Wages are measured as pesos uruguayos deflated by the official consumer price index (CPI). Firm size is measured as the log of total employment in each firm.

Table A2

Descriptive Statistics: Linked Employer–Employee Panel Data. Worker-level Information

	1997	2001	2005	2009
All workers employed in PCs				
Observations	9,634	9,533	10,265	12,706
Fraction female workers	0.31	0.36	0.41	0.45
Fraction members	0.40	0.42	0.45	0.41
Average age	41.08	42.59	42.83	41.88
Average job tenure	9.20	9.85	9.81	8.75
Gross monthly wage	25,538	23,675	17,154	19,355
Daily wage	982	1,004	679	805
Fraction in manufacturing	0.37	0.29	0.30	0.26
Fraction in transport	0.31	0.30	0.25	0.21
Fraction in services	0.30	0.39	0.42	0.48
Only those workers in WMFs				
Observations	3,270	3,202	3,898	4,417
Fraction female workers	0.15	0.14	0.24	0.27
Average age	42.23	44.02	44.61	43.94
Average job tenure	7.46	8.77	8.22	8.11
Gross monthly wage	23,757	22,594	16,243	17,629
Daily wage	944	890	666	811
Fraction in manufacturing	0.17	0.08	0.15	0.13
Fraction in transport	0.79	0.76	0.57	0.50
Fraction in services	0.04	0.13	0.22	0.28

Notes. Summary statistics are reported in October of each year. Wages are measured as pesos uruguayos deflated by the official consumer price index (CPI).

Table A3

Descriptive Statistics: Linked Employer–Employee Panel Data. Firm-level Information

	1997	2001	2005	2009
All PCs				
Number of firms	241	262	285	309
Firm size (log of employment)	2.69	2.57	2.63	2.63
Firm average wage	11,027	9,785	7,153	9,259
Coef. of variation (daily wages)	0.25	0.27	0.32	0.32
Fraction female workers	0.23	0.28	0.35	0.39
Average age	42.10	43.11	43.35	43.77
Age dispersion (SD)	9.63	9.47	9.57	9.84
Average job tenure	4.33	5.18	5.22	5.45
Job tenure dispersion (SD)	2.33	2.90	3.26	3.69
Fraction in manufacturing	0.25	0.18	0.19	0.18
Fraction in transport	0.44	0.40	0.33	0.26
Fraction in services	0.26	0.34	0.38	0.42
WMFs				
Number of firms	145	160	187	203
Firm size (log of employment)	2.50	2.37	2.52	2.54
Firm average wage	10,257	8,922	6,671	8,844
Coef. of variation (daily wages)	0.15	0.18	0.24	0.26
Fraction female workers	0.19	0.22	0.30	0.33
Average age	43.11	44.50	44.11	44.18
Age dispersion (SD)	9.50	9.44	9.53	9.74
Average job tenure	4.00	5.12	4.79	5.21
Job tenure dispersion (SD)	1.90	2.70	2.90	3.37
Fraction in manufacturing	0.25	0.19	0.19	0.20
Fraction in transport	0.59	0.53	0.40	0.31
Fraction in services	0.14	0.20	0.27	0.33

Notes. Summary statistics are reported in October of each year. Wages are measured as pesos uruguayos deflated by the official consumer price index (CPI). SD = standard deviation.

Table A4

Descriptive Statistics on Workers by Transition States

	% Female	Daily wage	Age	Tenure	Firm size	% Manufacturing	% Transport	% Services
Stayed in CFs	0.45	533.32	45.80	13.31	4.27	0.27	0.08	0.37
		(497.19)	(13.27)	(9.24)	(2.09)			
Stayed in WMFs	0.34	971.89	50.22	18.33	6.24	0.31	0.33	0.35
		(638.47)	(10.02)	(9.42)	(1.66)			
CF-to-CF movers	0.44	371.23	39.70	4.66	5.13	0.24	0.08	0.32
		(370.28)	(11.28)	(5.91)	(1.85)			
WMF-to-CF movers	0.50	672.30	46.18	10.71	6.06	0.16	0.13	0.60
		(553.70)	(10.66)	(8.99)	(1.65)			
CF-to-WMF movers	0.54	565.08	43.76	8.45	6.23	0.16	0.12	0.57
		(528.96)	(10.35)	(8.74)	(1.38)			

Notes. Wages are measured as pesos uruguayos deflated by the official consumer price index (CPI). Firm size is measured as the log of total employment in each firm. All variables measured at the time of entry. In the case of movers, tenure is measured at the maximum tenure reached previous to the first transition.

Table A5

Wage Gap Across the Wage Distribution. Results of Pooled Quantile Regressions. Period 1997–2009

	1997				2000			
	q20	q40	q60	q80	Q20	q40	q60	q80
Coop	0.175***	0.095***	0.021***	−0.033***	0.192***	0.107***	0.037***	−0.028***
	(0.006)	(0.005)	(0.006)	(0.007)	(0.007)	(0.005)	(0.006)	(0.007)
Age	0.037***	0.055***	0.066***	0.081***	0.044***	0.059***	0.068***	0.088***
	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)
Age^2	−0.001***	−0.001***	−0.001***	−0.001***	−0.001***	−0.001***	−0.001***	−0.001***
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
Female	−0.213***	−0.251***	−0.288***	−0.319***	−0.182***	−0.222***	−0.253***	−0.271***
	(0.003)	(0.002)	(0.003)	(0.003)	(0.003)	(0.003)	(0.003)	(0.003)
Tenure	0.056***	0.049***	0.045***	0.043***	0.063***	0.054***	0.0485***	0.045***
	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)
Tenure^2	−0.001***	−0.001***	−0.001***	−0.001***	−0.001***	−0.001***	−0.001***	−0.001***
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
Firm size	0.205***	0.183***	0.171***	0.166***	0.194***	0.173***	0.158***	0.149***
	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)
Test of interquantile differences								
20th = 40th	[0.000]				[0.000]			
20th = 80th	[0.000]				[0.000]			
40th = 80th		[0.000]				[0.000]		
Observations	389,190	389,190	389,190	389,190	389,055	389,055	389,055	389,055

Table A5
(*Continued*)

	2003				2006				2009			
	q20	Q40	q60	q80	Q20	q40	Q60	q80	q20	q40	q60	q80
Coop	0.142***	0.053***	−0.023***	−0.107***	0.159***	0.110***	0.040***	−0.040***	0.160***	0.114***	0.059***	−0.039***
	(0.006)	(0.006)	(0.006)	(0.009)	(0.007)	(0.005)	(0.007)	(0.009)	(0.008)	(0.005)	(0.005)	(0.006)
Age	0.039***	0.053***	0.064***	0.081***	0.039***	0.044***	0.054***	0.072***	0.031***	0.038***	0.045***	0.065***
	(0.001)	(0.001)	(0.001)	(0.002)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)
Age^2	−0.001***	−0.001***	−0.001***	−0.001***	−0.001***	−0.001***	−0.001***	−0.001***	−0.001***	−0.001***	−0.001***	−0.001***
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
Female	−0.153***	−0.195***	−0.239***	−0.277***	−0.167***	−0.185***	−0.226***	−0.264***	−0.202***	−0.211***	−0.240***	−0.263***
	(0.003)	(0.003)	(0.003)	(0.004)	(0.002)	(0.002)	(0.002)	(0.003)	(0.002)	(0.002)	(0.002)	(0.003)
Tenure	0.061***	0.060***	0.053***	0.050***	0.039***	0.042***	0.044***	0.043***	0.038***	0.044***	0.048***	0.049***
	(0.001)	(0.001)	(0.001)	(0.001)	(0.005)	(0.005)	(0.001)	(0.001)	(0.000)	(0.000)	(0.000)	(0.001)
Tenure^2	−0.001***	−0.001***	−0.001***	−0.001***	−0.001***	−0.001***	−0.001***	−0.001***	−0.001***	−0.001***	−0.001***	−0.001***
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
Firm size	0.188***	0.170***	0.159***	0.156***	0.125***	0.125***	0.129***	0.141***	0.106***	0.107***	0.112***	0.122***
	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)
Test of interquantile differences												
20th = 40th	[0.000]				[0.000]				[0.000]			
20th = 80th	[0.000]				[0.000]				[0.000]			
40th = 80th		[0.000]				[0.000]				[0.000]		
Observations	340,130	340,130	340,130	340,130	429,504	429,504	429,504	429,504	492,771	492,771	492,771	492,771

Notes. The dependent variable is the log of daily wages. The *Coop* dummy variable is set equal to 1 only for workers employed in a PC. Firm size is measured as the log of total employment in each firm. All estimates include six industry dummies. Bootstrapped standard errors (reported in parentheses) are based on 200 replications. ***Significant at 1%.

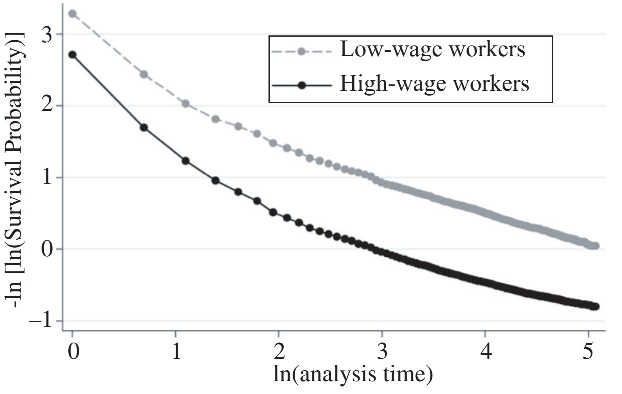

Fig. A1. *Graphical Check of the Proportional Hazard Assumption*
Notes. This graph plots the transformation $-\ln[-\ln\{\hat{S}(t)\}]$ *versus* $\ln(t)$ for high and low-wage members employed by WMFs, where $\hat{S}(t)$ is the Kaplan–Meier estimate of the survivor function. The proportional hazard assumption is not violated when the curves are parallel.

Table A6
Regression Analysis of Switchers Between Conventional Firms

	(1)	(2)	(3)
	Probit	Logit	Logit
Log Wage	−0.242***	−0.404***	
	(0.007)	(0.012)	
Wage Quintile 2			−0.035
			(0.022)
Wage Quintile 3			−0.172***
			(0.024)
Wage Quintile 4			−0.291***
			(0.026)
Wage Quintile 5			−0.333***
			(0.030)
Female	−0.122***	−0.201***	−0.152***
	(0.009)	(0.016)	(0.016)
Age	0.090***	0.152***	0.147***
	(0.004)	(0.006)	(0.006)
Age^2	−0.001***	−0.002***	−0.002***
	(4.94e-05)	(8.37e-05)	(8.34e-05)
Tenure	−0.133***	−0.220***	−0.225***
	(0.002)	(0.004)	(0.004)
Tenure^2	0.002***	0.004***	0.004***
	(0.0001)	(0.0002)	(0.0002)
Firm size	0.263***	0.443***	0.418***
	(0.003)	(0.005)	(0.004)
Observations	94,680	94,680	94,680
Pseudo R-squared	0.2052	0.2054	0.1982

Notes. The dependent variable is a dummy that takes a value of one if the worker experienced a CF-to-CF transition. Log Wage is the log of daily wage. Wage Quintile indicates the position of the worker in his sector-specific wage distribution. Firm size is measured as the log of total employment in each firm. All variables measured at the time of entry (before exit). In the case of movers, tenure is measured as the maximum tenure reached previous to the first transition. All estimates include a set of thirteen cohort dummies and six industry dummies. Standard errors reported in parentheses. *Significant at 10%; **significant at 5%; ***significant at 1%.

Table A7

Descriptive Statistics. Linked Employer–Employee 2009–12 Panel

	2009	2010	2011	2012
No. individuals	392,264	389,442	348,074	552,650
No. firms	45,455	49,910	44,960	76,132
% Female	0.405	0.416	0.423	0.435
Mean age	34.382	35.071	35.826	35.835
Mean annual wage	186,142	201,616	232,024	200,669
Mean log annual wage	12.134	12.214	12.355	12.209
% Manufacturing	0.322	0.311	0.276	0.283
% Transport	0.083	0.082	0.077	0.078
% Services	0.339	0.341	0.310	0.315
Mean firm size (in logs)	1.003	0.914	0.884	0.888
Mean firm wage (in logs)	11.140	11.165	11.306	11.234

Notes. Wages are measured as pesos uruguayos deflated by the official consumer price index (CPI). Firm size is measured as the log of total employment in each firm.

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