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The Mexican Wage Curve 2000–2003: A Quantile Analysis

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Abstract

This paper exploits the Mexican Encuesta Nacional de Empleo Urbano (ENEU) to determine the existence of the wage curve—an empirical phenomena first suggested by Blanchflower and Oswald (1990)—during the period 2000–2003. We propose an innovative approach to the wage curve by estimating the elasticity across the wage distribution. This is applied to the Mexican experience during the early 2000s recession. The evidence indicates that for Mexico during this period there is no wage curve, and that wages are positively affected by local levels of unemployment. This lends credibility to the Harris and Todaro (1970) view which suggests that there is segmentation in the labour market with residual unemployment. We argue that perhaps the power of unions may account for our findings.

JEL Classification: C21, J30, J60 O17

Keywords: Wage curve, unemployment, Formal/Informal Employment, Urban labour markets, Mexico

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1 Introduction

A recent strand of labour economics research has focused on the wage curve, the empirical phenomena which finds a negative relationship between unemployment and wages. This area of research was started by Blanchflower and Oswald (1990). They found that for the USA there was a strong negative elasticity with a magnitude of −0.10. Subsequent research across various countries has confirmed the existence of this relationship. Ramos et al. (2009, 2010) find that the wage curve elasticity in Colombia is −0.7 for all workers, however upon disaggregating this by the formal private and informal sectors, it becomes clear that the wage curve is primarily an informal phenomenon; manifested as highly negatively sloped wage curve. Nijkamp and Poot (2005) perform a meta–analysis of the wage curve literature, and determine that the curve on average can range from −0.5 to +0.1, excluding outliers. They find in favour of the Blanchflower–Oswald (B-O) assertion of the existence of a wage curve, however they estimate through a meta-regression that the 'true' elasticity of the wage curve is no more than −0.07. Berg and Contreras (2004) find that in Chile for the period 1957–73 there was no evidence of a wage curve. However, after economic reforms there is a wage curve with an elasticity of −0.08. They further disaggregate by employment sub-groups and find that the informal sector has not suffered lower wages as a result of unemployment, contradicting the view of the informal sector as a buffer during recessionary periods. Castro Lugo (2006) finds the wage curve to be −0.018 in the informal and −0.0099 in the private sector in Mexico during the period 1993–2002.

This 'empirical law' is rival to the models such as those proposed by Lewis (1954) and Harris and Todaro (1970). These models suggest that duality in the labour market is due to rural migration which creates an informal labour market, where informal workers seek higher non–agricultural wages in a formal or 'covered' sector of the labour market. Their failure in doing so creates a pool of residual unemployment. To minimise the effects of spells of unemployment individuals take up work in the informal sector whilst they queue for protected jobs. Another conclusion which can be drawn from this strand of research is that where there is a high level of demand for rationed jobs there will be frictional unemployment. Thus, we may expect there to be higher levels of wages as a reward and incentive for workers to remain in the formal sector. This suggests the existence of a positive relationship between wages and unemployment (Blanchflower and Oswald, 2005). Therefore, we can stylise this view as one where there are two distinct segmented labour markets (formal and rural) whilst the informal market acts a buffer to unemployment whilst workers wait for jobs in the formal sector.

There is some evidence which suggests that neither view fully characterises the labour market. Papps (2001) finds that the wage curve in New Zealand in the short run is −0.98, however for the long–run levels of unemployment he finds that the wage curve is positive thus providing some evidence for the Harris–Todaro (H-T) view of a positive relationship between wages and unemployment. Partridge and Rickman (1997) employ a different sample than that of B-O and conclude that the wage curve for the USA may not be an empirical law. They find evidence for the Harris and Todaro view of the relationship between wages and unemployment. They further comment that their evidence for the wage curve's existence is more akin to a Phillips curve than what is proposed by B-O. Castro Lugo (2006) finds that in Mexico there is no evidence of a wage curve in the public sector.

There has been a large effort in investigating the effects of the debt and currency

crises on the Mexican economy. However, as Bosch and Manacorda (2008) point out very little research has analysed the specificities and particularities of the Mexican labour market, for the period post–1994 and in particular the to the 2003 recession. This period is particularly interesting as in 1994, Mexico became a signatory to the North American Free Trade Agreement (NAFTA). This led to the Mexican economy becoming heavily dependent on the state of the American economy. By 2000, 88.7% of exports and 73.1% of imports were to and from the USA respectively. One of the industries which benefited from NAFTA was the *maquila*, or manufacturing sector. This intermediate industry alone accounted for 47.7% of Mexican exports and 34.4% of imports in 2000. Due to the nature of this industry it is heavily entwined with the rate of industrial activity in the USA. The fall in industrial production in the American economy had severe consequences for the Mexican labour market. From October 2000 to March 2002 there was a fall in *maquila* employment of 21%, and a fall in production of 8%. These decreases in production and employment, whilst related to the Mexican dependence upon the American economy, are by no means the sole factors in the contraction of the manufacturing sector. Robertson (2000) tests whether shocks to wages in the USA affects Mexican wages, he finds that an increase of 10% in wages north of the border increase wages in urban Mexico by 1.8%. These findings suggest that just as output in Mexico has a strong correlation with US industrial output, wages also have a similar relationship. The lack of labour market and enterprise reforms during this period led to Mexican competitiveness being eroded, a large portion of this manufacturing transferred from Mexico to Asia (Hanson and Harrison, 1999).

An important aspect in the Mexican labour market, which we thus far have not considered is the role the informal sector. Throughout the literature there is very little consensus on an exact definition (Pradhan and van Soest, 1995; Gasparini and Tornarolli, 2007; Loayza and Sugawara, 2009) of informality. The multitude of definitions can be synthesised into two competing views: The 'Productive' view of informality, where the informal sector is characterised by "workers in low-productivity jobs in marginal small–scale and often family based activities" (ILO, 1991). Under this definition the formal sector is likely to be rationed in the number of people it can employ and as such informality is a by–product of waiting for a covered sector job. The second and much more recent definition is recognised by the ILO (2002) and focuses on the legalistic approach, wherein informality is classified as being in a state by which one's work is "not recognised or protected under the law and therefore receive little or no legal or social protection and are unable to enforce contracts or have security of property rights... [and] are excluded from or have limited access to public infrastructure and benefits." This definition accepts more broadly that informality is not a simple concept to classify, and as such a means by which we may observe it is as a residual activity: Those whom are not able to avail themselves of the labour laws.

Within the H-T view it is possible to characterise the informal market as a symmetric and competitive alternative to the formal sector. This view is proposed by Heckman and Sedlacek (1985) who postulate that there are multiple sectors, and workers will self select into the sector that maximises their expected wages. They apply the well known Heckman (1979) correction procedure and proceed to estimate the determinants of wages for the USA. This type of methodology is exploited by Magnac (1991) who applies it to Colombia and finds that there is a competitive two sector labour market. Pradhan and van Soest (1995) find that in Bolivia there is labour market segmentation and that the

informal market is highly competitive vis-a-vis the formal market. Maloney (1999) analyses whether the presence of informality is indicative of segmentation in the labour market in Mexico. Utilising worker transitions between sectors he finds that the informal labour market is attractive to workers within itself, and not a place to wait for formal sector jobs. Gong and van Soest (2002) find that wages in both sectors increase with higher levels of education, and high levels of mobility. They also find that for the lower educated workers the dualistic view is not a good descriptor of the labour market. However, for the highly educated there would appear to be strong rigidities which give rise to dualism.

Gasparini and Tornarolli (2007) find that informality in Latin America follows no common trend across all countries. However, they find that increases in informality in Mexico are pro-cyclical. During economic expansions they find that decreases in informality are coupled with increases in informal wages: the type of behaviour which would be expected in a competitive market. These transitions are confirmed by Bosch and Maloney (2007). They find that informality, whilst decreasing, also sees a considerable amount of flows into self–employed informality counter-cyclically. Their findings also suggest that the informal sector itself may be heterogeneous in nature, as there seems to be voluntary informality, but at the same time a sizeable section which corresponds to the traditional H-T view. They find that this is particularly the case for young workers. Bosch and Maloney (2006) find, interestingly, that during the 1995 crisis in Mexico the changes in unemployment were driven by destruction of informal jobs rather than any effects upon the formal sector. They also find that for the 2001 recession, the main driver in terms of cyclical effects was the destruction of formal sector jobs, since the industry primarily affected was export manufacturing, which is predominantly composed of formal sector employees.

This brief review of the existing literature has motivated the question whether a wage curve exists for Mexico. As we have seen there are two rival theories of the relationship between the rate of unemployment and wages. The Harris and Todaro (1970) view suggests that there will be a positive relation, whilst the more commonly accepted view proposed by Blanchflower and Oswald (1994) suggests the opposite. Determining which relation characterises the Mexican labour market is integral for understanding the wage determination processes in the Mexican economy, and how they are affected by cyclical flows in unemployment.

This paper will focus on the Mexican urban labour market. In particular, we aim to determine the existence of the wage curve and whether it was affected by the recession of the early 2000s. We will further disaggregate into the three sectors of the economy: the informal, private and public sectors. We adopt a legalistic approach to informality, which allows us to identify informal workers. We utilise the Encuesta Nacional de Empleo Urbano (ENEU)—a household survey of the labour force—to perform our analysis. We do so by applying the traditional wage determination approach of Mincer (1974). This is augmented by selection correction methods proposed by Lee (1983). We have identified a gap in the literature which we shall attempt to expand on: whether the wage curve changes throughout the wage distribution. To this aim we shall estimate it utilising quantile methods (Koenker and Bassett, 1978; Koenker, 2005).

This paper is organised in the following way. In section 2 we outline the limitations and advantages of the data. Section 3 outlines the econometric methodology of the study. Section 4 reports the results, and section 5 concludes with a discussion the findings.

2 Data

The data used in this analysis are drawn from the Mexican National Urban Employment Survey (ENEU). This survey is carried out at the household level and was carried out continuously from 1984–2004 quarterly, and was conducted by the Mexican Institute of Statistics and Geography (INEGI). Geographically this survey comprises all the states in Mexico, with the main cities of each being included. We selected the third quarter of 2000 and 2003 as it would give us an adequate time-period before the recession and one time-period at the tail end of the recession thus allowing for a meaningful comparison to be made of the impact of the recession.

The number of households surveyed was determined by a series of factors such as the average number of inhabitants being 3.14. The net economic participation being between 48-52%[2]. Therefore the sample size of households was determined to be 2,100 for most cities with the exceptions of Mexico City (5,100), Guadalajara, Monterrey, Puebla, Leon, and Torreon, Mexicali (3,000), La Paz and Cancun (1,800).

The design of this survey is constructed probabilistically, stratified, conglomerated and through three phases. It is probabilistic as the probability of inclusion of a unit is known and different from zero for each member of the population. Triple phased since the primary units of survey are first identified though AGEB[3], followed by the secondary units which organised among block lines where there is a minimum of 40 inhabited households, and finally the inclusion of non-permanent households[4] are selected into the sample. Stratification is carried out through socio–economic factors in a multivariate framework. The survey selection is conglomerated as a unit of survey was determined by homogeneous housing and heterogeneous inhabitants.

For the purposes of our analysis we restricted the sample to all males aged between 16 to 65. Females were not selected as it was deemed that in order to correct the well known selection bias into the labour market as proposed by Heckman (1979), there would be insufficient instruments to correctly identify the selection effects and thus complicating the use of this procedure. The sample utilised for calculating the selectivity bias of the work sectors was inclusive of those individuals who are unemployed, self employed and out of the labour force. The main analysis restricted the sample further to all salaried individuals, regardless of their work status in the reference week. We concentrate on main job earnings and ignore any secondary jobs to minimise the introduction of reporting and measurement errors that may be attached to secondary earnings. The decision to not control for self-employed individuals was based on the possibility that they would be subject to different wage determination and selection processes. The quality of the available instruments for identification was not sufficient for their inclusion to be justified.

We adopt the definition of informality promoted by ILO (2002), we define informal workers as those who report themselves as having a contract but do not have access to social security benefits through the *Mexican Institute of Social Security* (IMSS). The data allows for this as there is a comprehensive question on benefits received from work. The identification of public sector workers was done in a similar manner, workers who indicated they were covered by the *Institute for Social Security and Services for State*

[2]This rate was determined by previous surveys including the national census.

[3]Each of this is divided along political lines into state, municipality and finally into a statistical construct by INEGI called Basic Geo-statistic Area (AGEB) which is an area of roughly equal size comprising 480 households.

[4]Such as those living in train coaches, trailers, boats, caves, etc.

Workers (ISSSTE) were determined to be public sector workers.

The data allow us to distinguish the educational qualifications an individual has obtained, the age of the respondent, occupation, the region of their residence, and hours worked. The survey allows individuals to report their wages as they prefer, be that weekly, daily, monthly, annually as well as per unit. However, the dataset does allow for this to be transformed to an equivalent net weekly wage. The wages were deflated by the GDP implicit price deflator obtained from the UN Statistics Agency and are expressed in 1990 constant prices. Whilst regional deflators might be a desirable addition we were unable to obtain them for this sample period.

The decision to focus on ENEU was due to the periodicity of the survey. A survey which covers both urban and rural areas of Mexico was run by INEGI but it was performed yearly. We had a trade–off between quarterly data and yearly data, as the yearly data may hide the fluctuations in income due to an exogenous negative shock. Therefore the decision was made that due to the nature of the study it would be preferable to focus on the urban areas where majority of the informal and public sector activity is likely to be found.

3 Econometric Methodology

Following in the tradition of labour economics in estimating empirically the wage relationship as in Mincer's (1974) seminal work, we shall estimate log wages as a function of wage determining characteristics. The specification will include variables capturing human capital controls and other variables deemed important in the wage determination process. In the context of the wage wage curve we expand our specification according to Blanchflower and Oswald (1994). This gives the sector specific labour market earnings equations for the i^{th} individual in the j^{th} and l^{th} sectors in the r^{th} region are given as follows,

$$w_{irj} = \ln u'_{rj}\lambda + x'_{irj}\beta_{irj} + v_{irj} \tag{1}$$

$$w_{irl} = \ln u'_{rl}\lambda + x'_{irl}\beta_{irj} + v_{irl} \tag{2}$$

where $\ln u_{rj}$ and $\ln u_{rl}$ are a (n $\times$ 1) vector of the natural log of regional unemployment, x_j and x_l are (k $\times$ n) matrices of characteristics (e.g. education, type of contract, employment industry, regional controls, number of coworkers, etc.) and β is a (k $\times$ 1) vector of unknown parameters which capture the effect of the various covariates on the natural log of the wage (w), v is a (n $\times$ 1) vector of random error terms specific to each sector.

We might expect that there is an element of self–selection into the different sectors of the economy. We shall now motivate a selection correction mechanism. Assuming that the utility function for the i^{th} individual in sector z with benefits associated with that sector b has a utility function of the following form:

$$U_i(z,b) = \alpha \cdot w_{iz} + \kappa'_i\gamma_z + \eta_i(z,b) \tag{3}$$

where w is the natural log of hourly wages and γ is a vector of individual characteristics. It follows that an increase in w would increase the individuals utility. An individual will prefer to be in sector z if the following is true:

$$U_i(z,b) + C(z,b) \geq U_i(l,b) + C(l,b) \tag{4}$$

That is sector z with associated benefits b will be preferred as long as the value added C to the individual in sector z with associated benefits b are greater or equal to that in sector l and its associated benefits b. This implies that the distribution of individuals across the sectors may not necessarily be random. Therefore, estimation must include a means by which to control selection prior to any wage equation estimation. The exact value of the value added of the benefits associated with sector z, $C(z, b)$ is unobservable in (3). In the empirical application of this we shall employ a variable for the number of household members and a dummy controlling for household head status, and shall assume that these will capture all of the added benefits of being in sector z.

This formulation has some caveats associated with it. As it is based on a strong set of neoclassical assumptions this ignores employee sample selection, and barrier to entry. It is also particularly problematic as the marginalisation of individuals who may be discriminated against despite having all of the necessary qualifications but unable to break into the formal or public sector employees. However, whilst it is an imperfect way to explain the motivation for an individual to be in a particular sector we shall employ this to motivate the rest of the methodology.

3.1 Selection Methodology

The foundation for this type of strategy was set out by Heckman (1979) within a bivariate framework. He suggested two methods by which to correct for issues related to selection: the Full Information Maximum Likelihood Estimation (FIML) and a two step estimation, with the first being a selection model utilising a participation equation from which a selection correction term, also called the inverse Mills ratio, can be calculated. This is included in a simple regression equation estimated by Ordinary Least Squares (OLS). Due to the intended outcomes being multinomial and not binary we shall adopt the approach suggested by Lee (1983) which is a generalisation of the Heckman procedure to a multivariate case.

Substituting a sector and region independent[5] version of (1) and (2) into (3) gives us utility in a reduced form, with sector and region specific subscripts suppressed:

$$U_i(z, b) = \alpha \cdot (x'_{iz}\beta_z) + \kappa'_i\gamma_z + \epsilon_i(z, b) \tag{5}$$

where $\epsilon_i(z) = \alpha \cdot u_i + \eta(z)$.

Assuming that $\eta_i(z)$ has a type–I extreme value distribution it can be shown to be:

$$Z(i \text{ in } j^{th}\text{sector}) = \frac{exp(x'_{ij}\beta_j^{\alpha} + \kappa'_i\gamma j)}{\sum_{z=j}^{Z} exp(x'_{ij}\beta_j^{\alpha} + \kappa'_i\gamma_j)} \tag{6}$$

where $Z = j, \ldots, l$ is the total number of labour statuses and $\beta_j^{\alpha} = \alpha \cdot \beta_j$.

This is the multinomial logit (MNL) model as developed by McFadden (1973). In order for estimation to provide unique estimates a Theil normalisation must be performed on an arbitrary category so that the model may be identified. The κ vector therefore must include all the exogenous variables of equations (1) and (2) as well as identifying instruments which allow for the identification of the choice which an individual takes when

[5] For simplicity we conflate the term $\ln u'\lambda$ with the vector $x'\beta$

selecting a labour market option to enter. Therefore we shall employ this methodology utilising the estimates from a five category MNL for the various labour market outcomes[6].

It should be noted that the Multinomial Logit does possess a flaw which must be considered when utilising it, namely the Independence of Irrelevant Alternatives (IIA) assumption which for this case would state that $cov(\epsilon_{(l=1)}, \epsilon_{(l=2)}) = 0$ however since the error term is derived from the error terms of equations (1) (2) and (3) we can expect that this might be violated. However, based upon Monte Carlo simulations Bourguignon et al. (2007) have shown that the violation of IIA should not be a major concern "when the focus is on estimating an outcome over selected populations rather than on estimating the selection process itself". Therefore giving us confidence that the results will not be biased in the face of IIA violations. Therefore as shown by Lee (1983) we can correct for endogenous selection into labour market outcomes in the earnings equations through adapting the Heckman two–step procedure to the case of polychotomous choice models. One last caveat that must be acknowledged is that by employing the Lee (1983) methodology we are making strong assumptions about the structure of the correlation between disturbances of the selection equation and earnings equation as demonstrated by Bourguignon et al. (2007). These assumptions suggest that the correlation between the error term of the earnings equation (u_i) is uncorrelated with the conditional joint distribution of the error terms in the selection equation (ϵ_i). This is likely to be violated, although the errors are transformed due to the multinomial logit this only ensures that the marginal distribution of errors are independent it is unlikely the joint distribution of errors from the selection equation and earnings equation are independent. The implications of this violation are that the correlations of the polychotomous choices and the unobserved determinants of income are identical across all choices. Whilst this may be a strong assumption we believe that the trade-offs when compared to the other alternative selection mechanisms such as the Dahl (2002) and Dubin and McFadden (1984) are justified under the same condition as with the IIA violation, that is we are interested in the wage process in a selected population of wage earners and not the selection process in itself. Thus we adopt the selection correction term as suggested by Lee (1983) is obtained by implementing the following:

$$SCT_j = \frac{\phi(\Phi^{-1}(P_j))}{P_j} \tag{7}$$

Where P_j is the estimated probability from the multinomial logit for the jth outcome $\phi(.)$ is the PDF of the normal distribution and Φ^{-1} is the inverse of the CDF of the normal distribution. This yields a selection term for each of the outcomes of the dependent variable. The selection term for a given outcome is introduced to correct for unobservables in the selection process in the mean regression analysis.

In terms of selection effects and the quantile regression, there are methods such as Buchinsky (1998) who employs higher order expansions of the inverse Mills ratio. These unfortunately complicate the identification of the constant term. We opt to use the mean selection correction term. We acknowledge that it will not completely correct for selection, but it is believed that it will correct some of the selection effects in a satisfactory manner.

[6]Namely: Out of the Labour force, Unemployed, Informal, Formal Private and Formal Public.

Table 1 OLS Estimates of wage curve for Mexico, 2000 and 2003

	2000				2003			
	Pooled	Informal	Private	Public	Pooled	Informal	Private	Public
Wage Curve	0.062***	0.078***	0.051***	0.024	0.120***	0.150***	0.093***	0.079**
	(.0052)	(.0082)	(.0077)	(.0150)	(.0092)	(.0147)	(.0128)	(.0297)
State Controls	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Ind. Controls	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
R^2	0.4046	0.3423	0.4006	0.3922	0.3955	0.3201	0.3863	0.3862
N	59547	26131	28117	5299	39605	18798	17299	3508

Notes: ***,**,* denote significance at 1, 5 and 10 per cent respectively.
The standard errors reported in parentheses were calculated using the White(1980) sandwich estimator. The pay measure is net of taxes and is measured as the log of the hourly wage in constant 1990 prices. The regressions included 12 industry controls, and 31 regional controls. Selection terms are included for all regressions.
Full results can be found in table A2 in the appendix

4 Results

4.1 Multinomial Logit Selection Regression

As shown in the previous section estimates of the determinants of wage relationships could be potentially biased by selection. Therefore prior to estimating any wage equation we want to control for individuals selecting into the different sectors of the labour market. To do so we must estimate a multiple outcome model. The underlying approach to the MNL estimation and the estimates can be found in table A1 of the appendix. Note that these are not interpreted as they did not pass the Small–Hsiao IIA test. The failure of IIA does not concern us in this instance, this is due to evidence by Bourguignon et al. (2007) which suggests that although IIA may be violated selection terms based on an MNL model are not biased.

4.2 Mean Analysis

The mean regression includes variables for age and its quadratic. This was chosen over potential experience due to the possibility of an individual starting to work prior to the end of compulsory education at age 16. We include a dummy to capture the effects of marriage on earnings. We also include a set of dummies capturing maximum educational qualification obtained. These were preferred over a continuous variable capturing the number of years in education, and the rationale behind this decision was two–fold. Firstly this allows for the coefficients to differ based on the level of education, whilst at the same time allowing ease of use for computation of returns to education. Secondly the dummy set is less restrictive in terms of making assumptions on the number of years an individual takes to complete a qualification; the structure of the survey questions do not easily translate into a continuous variable and as such may have been subject to measurement error. The augmented Mincerian equation specification includes thirteen industry controls, dummies capturing the type of contracts available to an individual, as well as controls for all 32 Mexican states. For the pooled regression, dummies to capture the public sector pay gap and the informal pay gap were also introduced. We also included the state level unemployment rate for the relevant quarter and year. This will be utilised to determine the whether a wage curve exists for Mexico as suggested by Castro Lugo (2006) and originally proposed by Blanchflower and Oswald (1990). We report the augmented specification in Table 1.

Table 1 reports the estimates of the Mexican wage curve for 2000 and 2003. It should

be noted that selection effects for the informal sector in 2000 as well as the private sector in both years. In terms of the wage curve our findings are at odds with the existing literature. As mentioned earlier Blanchflower and Oswald (1995) find a wage curve elasticity of about −0.10, whilst Ramos et al. (2009, 2010) find an elasticity of −0.07 for Colombia, and Castro Lugo (2006) finds an elasticity of −0.0099 for the formal sector and −0.018 for informal sector in Mexico. The estimates of the 'wage curve' elasticity at the mean we find all range between 0.024–0.150. These are all well determined, and the between-sector difference in elasticity is found to be statistically significant, with the sole exception of the difference between the public and private sector in 2003.

Our results show that for all sectors at the mean, higher levels of unemployment imply higher wages. This is particularly strong for the private sector in 2000, the differential of 0.038 is found to be statistically significant from the pooled estimate. The estimates imply that a 10% increase in local unemployment increases public sector wages by 5%. We argue that on the basis of these findings this is suggestive that at the mean there would certainly appear to not be a wage curve. The absence of a wage curve lends evidence to the Harris and Todaro (1970) view wherein the higher wages are a means to compensate for the prospect of unemployment.

4.3 Quantile Analysis

We now turn our attention to the quantile regression estimates. The specified wage equation included: age and its quadratic; marital status; variables controlling for highest qualification achieved; variables controlling for contract types amongst workers; the state level local unemployment rate for each worker; industry controls; state controls; a selection term where appropriate. The regressions were run with the selection terms; if they were found to not be significant we omitted them and reran the regressions. The estimates we report are from the second set of regressions. The omission of selection effects was intentional as we tried to avoid biasing the estimates by the introduction of selection effects where none were relevant.

The quantile estimates will allow for the coefficients to vary across the wage distribution. It should allow us to see whether the estimates we found at the mean are applicable throughout the whole of the wage distribution. This approach is innovative in terms of the wage curve literature. To the best of our knowledge there have not been quantile estimations of the elasticity of the relationship between wages and unemployment rates.

We shall first examine the median regression estimates. Table 2 reports the pooled and sector specific estimates of the wage curve; overall the fit of the models on the sub–samples is poorer than for the mean regression estimates. However, for the individuals in the informal and private private sector the estimates are very well determined. In terms of the public sector the estimates are found to not be as consistent in terms of significance.

The existing literature has thus far neglected the changes in elasticity across sectors and along the wage distribution. The evidence we report for the median wage curve elasticity estimates is similar to that at the mean, but consistently more elastic. We find that the elasticity is characterised by the Harris & Todaro view, wherein wages and unemployment have a positive relation. In 2000, for the whole of the labour market we find a wage curve elasticity of 0.056, this implies that a 10% increase in local unemployment increases wages by 5.6%. This effect is significantly different for all of the labour market sectors.

The informal sector is found to have a positive wage curve elasticity. We find that it is significantly different from the pooled estimate. This estimate is also found to have

Table 2 Quantile Regression Estimates for Wage Curve at the median, 2000 and 2003

	2000				2003			
	Pooled	Informal	Private	Public	Pooled	Informal	Private	Public
Wage Curve	(.2447)	0.073***	0.049***	0.007	0.106***	0.147***	0.077***	0.055*
	(.0050)	(.0088)	(.0080)	(.0153)	(.0084)	(.0129)	(.0128)	(.0337)
State Controls	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Ind. Controls	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Pseudo–R^2	0.2587	0.2151	0.2428	0.2503	0.2484	0.1927	0.2344	0.2514
N	59547	26131	28117	5299	39605	18798	17299	3508

Notes: ***,**,* denote significance at 1, 5 and 10 per cent respectively.
The pay measure is net of taxes and is measured as the log of the hourly wage in constant 1990 prices. The regressions included 12 industry controls, and 31 regional controls.
Selection terms are included for Private sector in 2000 & 2003; and Informal, 2003
Full estimates can be found in table A3 in the appendix

increased significantly over the recession. Thus in 2000, a 10% increase in unemployment increased informal sector employees earnings potential by 7.5%; this increased to 15% by the end of the recession. We may be able to account for this increase as being a result of the informal sector being utilised as a wait status to join the private sector. This would be consistent with the evidence reported for short term contracts. It would appear that trial periods for new workers under Mexican labour laws may have been exploited by employers in order to decrease their payroll contributions to social security.

In the private sector there is a highly significant difference between the wage curve elasticity reported and that at the median. The elasticity is found to significantly increase over the time period surveyed by 0.027, with an absolute t-ratio of 1.82. Thus, over the recession a 10% increase in state level unemployment increased wages earnings potential for private sector employees by 2.7%.

For the public sector there is no wage curve in 2000. In 2003, there is a significantly positive relationship between public wages and unemployment at the median, perhaps signalling that there is some level of queuing for these higher paid jobs. The effect suggests that in the public sector a 10% increase in unemployment leads to 8% higher wages.

Selection effects are present for the private sector in both years and they are found in the informal sub–sample by the end of the recession. These effects suggest that the median individual who selects into the private sector earned 20.6% more than a random selection of individuals. The selection effect whilst still being present seems to have decreased after the recession to 15%. The median individual who selected into the informal market in 2003 earned 8.2% more than a selection of random individuals. The presence of selection effects into the informal market once again suggest that perhaps the informal market is utilised as an escape valve in the face of adverse economic conditions.

The estimates at the 10th quantile are reported in Table 3. These are less well determined than those at the median. What becomes slightly more troublesome is the poor determination of the public sector, albeit it curiously has the highest pseudo–R^2 of all of the sub–samples, particularly in 2003.

We should note that no effects seem to be present in terms of contracts for the informal sector, suggesting that the cushioning role of the informal labour market only has effect on the upper ends of the wage distribution.

Selection effects in the informal sector in 2000 are found to be significant: individuals who selected into the informal sector in this sample have an earnings potential 10% higher than a random sample. For the private sector we find that there are selection effects for 2000 and 2003. These imply that individuals who selected into the informal market in

Table 3 Quantile Regression Estimates for Wage Curve at the 10th quantile, 2000 and 2003

	2000				2003			
	Pooled	Informal	Private	Public	Pooled	Informal	Private	Public
Wage Curve	0.053***	0.071***	0.020**	0.024	0.115***	0.115***	0.109***	0.010
	(.0067)	(.0113)	(.0100)	(.0196)	(.0134)	(.0220)	(.0167)	(.0448)
State Controls	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Ind. Controls	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Pseudo–R^2								
N	59547	26131	28117	5299	39605	18798	17299	3508

Notes: ***,**,* denote significance at 1, 5 and 10 per cent respectively.
The pay measure is net of taxes and is measured as the log of the hourly wage in constant 1990 prices. The regressions included 12 industry controls, and 31 regional controls.
Selection terms are included for Private sector in 2000 & 2003; and Informal, 2000
Full estimates can be found in table A4 in the appendix

2000 earned 15.8% higher wages than a random sample. In 2003, the selection effect was found to be 13.5%

We now explore the wage curve evidence. The elasticities estimated for the wage curve are found to vary significantly across sectors. Further interrogation of the data reveals that there are statistically significant changes in this variable over the recession for the pooled, informal and private estimates.

The informal sector, at the bottom end of the distribution in 2000 is found to have an wage curve elasticity of 0.71, this increased by 2003 to 0.115. This implies that a 10% increase in unemployment in 2000 would increase informal sector earnings potential by 7.1%. At the end of the recession this was found to have increased to 12.2%. There is a similar story in the private sector. In 2000, a 10% increase in unemployment would increase wages by 2%. By 2003, a shock of the same magnitude would increase wages by 10.9% It would appear that there is no relationship for the public sector at the bottom end of the distribution: this suggests that public sector wage determination is independent of the levels of unemployment.

At the 90th quantile we find that the estimates are well determined for 2000, and have much higher pseudo–R^2 than at the 10th quantile. The 2003 estimates in terms of the contract variables are poorly determined. The pseudo–R^2 for these estimates is lower than the 2000 sample.

In general we find that there are sector specific wage curve elasticities, with the notable exceptions of the private sector in 2000 and the informal sector in 2003. We find that the pooled, informal and public estimates of the elasticities vary significantly over the recession, with t-ratios of 3.12, 2.16 and 2.41 respectively. These changes are all relative increases in elasticity. It is salient that the unemployment elasticity for the public sector in 2003 is 0.21: this implies that there is a positive trade–off between higher wages and unemployment, where a 10 percent increase in local unemployment increases wages by 21%. The difference of this effect with respect to the private sector is found to be 12% with statistical significance and an absolute t-ratio of 2.54. Whilst there might be wage increases *vis-a-vis* unemployment across all sectors the magnitude of this effect is by far the largest.

One mechanism which may explain the positive wage curve elasticities that we report across the quantiles may be the role of unions on wage setting in Mexico. This would be similar to an insider/outsider model of the labour market, wherein the insiders or union members seek to increase wages at the expense of higher levels of unemployment. This distortion would therefore be reflected by a positive elasticity. The magnitude of the elas-

Table 4 Quantile Regression Estimates for Wage Curve at the 90th quantile, 2000 and 2003

	2000				2003			
	Pooled	Informal	Private	Public	Pooled	Informal	Private	Public
Local Unemp.	0.072***	0.099***	0.056***	0.038*	0.136***	0.160***	0.082***	0.211***
	(.0099)	(.0149)	(.0141)	(.0230)	(.0179)	(.0233)	(.0230)	(.0450)
State Controls	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Ind. Controls	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Pseudo–R^2								
N	59547	26131	28117	5299	39605	18798	17299	3508

Notes: ***,**,* denote significance at 1, 5 and 10 per cent respectively.
The pay measure is net of taxes and is measured as the log of the hourly wage in constant 1990 prices. The regressions included 12 industry controls, and 31 regional controls.
Selection terms are included for Private sector in 2000 & 2003.
Full estimates can be found in table A5 in the appendix

Table 5 90th–10th Quantile Estimates for the Wage Curve

Pooled		Informal		Private		Public	
2000	2003	2000	2003	2000	2003	2000	2003
0.0195*	0.0214	0.0319*	0.0121	0.0003	-0.0225	0.0143	0.201***
(0.0112)	(0.0189)	(0.0185)	(0.0345)	(0.0155)	(0.0240)	(0.0263)	(0.0705)

Notes: Specifications are the same as tables A2–A4 in appendix
***,**,* denote significance at 1, 5 and 10 per cent significance.
Standard errors reported in parentheses

ticity might also indicate the relative wage bargaining power of the union.

To check the robustness of the wage curve estimates we report, we performed inter–quantile estimates for the pooled, informal, private, and public sector. These are reported in Table 5. The estimated differences between the 90th and the 10th quantile are generally found to be positive. Testing the differences in estimates with the pooled model we find that there are no differences in 2000. For 2003, we find that the pooled public differential is statistically significant with an absolute t-ratio of 2.41. In terms of temporal differences over the recession, we find that the public sector is found to be significant with a t-statistic of 2.47. These findings confirm the positive effect which we report in 2003 for the public sector, where a 1% increase in local unemployment will lead to differential in increases of wages of 2% between the top and bottom ends of the wage distribution.

5 Discussion

In this paper we set out to investigate the existence of a wage curve in the Mexican labour market during period 2000–2003, and thus fill the gap in the literature, particularly with respect to Mexico by estimating the wage curve as first suggested by Blanchflower and Oswald (1990) . This is of particular interest as there is very little coverage in terms of empirical analyses; to the best of our knowledge there is no comprehensive analysis of this kind for the period which we surveyed. We also innovated on the existing literature by applying quantile methods to see whether this varies across the wage distribution.

Our main findings suggest that there is no wage curve present in Mexico for the years 2000-2003, instead it we find that unemployment and wages follow the more traditional Harris and Todaro (1970) view, wherein higher levels of unemployment lead to higher wages. For the informal and private sector the elasticities are found to increase monotonically through the wage distribution, having the most elastic effect at the upper ends of the distribution.

We should note that the public sector is not always sensitive to local labour market conditions. In general, for 2000 we find no wage curve relationship at the mean or median, but it becomes significant at the upper end of the wage distribution, suggesting that at the mean and median w age determination is independent of local labour market conditions. In 2003, we find that there is a significantly positive relationship for the mean, median and 90th quantile. The estimates we obtain for the wage curve elasticity are found to have significantly become more elastic over the recession. To invee stigate whether the wage curve differs across the wage distribution we performed inter-quantile estimates. We found that for the pooled, informal and private model there are positive differences between the 90th and the 10th quantiles. Further interrogation suggests that the private sector follows the same wage curve distribution for the pooled model.

The findings we report are at odds with the wage curve literature (Blanchflower and Oswald, 1990, 1994, 2005; Ramos et al., 2009, 2010; García-Mainar and Montuenga-Gómez, 2003; Papps, 2001). In particular our findings are contrary to those of Castro Lugo (2006) who finds the existence of a wage curve for the period 1993–2002. We argue that the robustness of our findings are explained since the positive effect we find is consistent across and in between all the quantiles, as well as at the mean.

We argue that the 'empirical law' of the wage curve is nullified due to the rise of independent unionisation in Mexico. In general the evidence from the industrial relations literature suggests that unions have three effects upon the dispersion of wages (Fairris, 2003). They decrease the between union worker dispersion of wages, they also decrease the union-nonunion wage differentials. However, this comes at the cost of increasing the dispersion between union individuals and those who they are unable to unionise. Putting this into the perspective of the findings we report we would expect the effects of unions to decrease the wage dispersions of the unionised industries, as well as the public private wage gaps, whilst increasing the wage gaps between the unionised sectors and the informal individuals.

In order to better understand the why the rise of independent unions may affect the wage curve it is useful to understand the history of the Mexican union movement. Middlebrook (1995) explains how during the founding of the *Partido Institucional Revolucionario*[7] (PRI) the largest union in Mexico the *Confederación de Trabajadores de México*[8] (CTM) there was a power sharing agreement where they received state money in exchange for allowing the government errode the bargaining process of workers and exert control over labour relations. Although this union did not hold a monopoly on organising labour they ensured that independent unions were marginalised. However, the role of the 'official' union was strongly tied to the PRIs fate. The CTM's power began to decline in the 70s. By the early 80s independent unions began to organise labour but remained irrelevant. However the 90s saw a period of privatisation and further erosion of the power of the CTM, this coupled with the death of their leader Fidel Velázques in 1997 led to an inexorable decline of the official unions. This is picked up by Fairris (2003) who finds that unionisation in Mexico has decreased since the late 1980s and this has led to increases in wage inequality between sectors. However, this decline of union power has not occurred symetrically across all sectors of the economy, for example majority of the public sector remains unionised as workers are required to join 'official' government unions. During

[7]Political party which held a monopoly on political power from its founding in the 1930s until it was ousted in the 2000 elections

[8]Confederation of Mexican workers

this time period old unions such as the teachers union (SNTE) and new ones such as the national union of workers (UNT) have taken an increasingly tough approach to wage bargaining which means their wage bargaining power has not decreased equiproportionately to the decline in membership of the whole spectrum of unions(Fairris and Levine, 2004). This may be indicative that although there was a marked decline of unionisation rates during this period there has been a structural change in the composition of unions and an increase in their effectiveness.

We argue that the rise of independent unions which conform to the stylised facts suggested by the empirical literature on industrial relations may account for the lack of a wage curve during this period. This is perhaps quite a novel idea as there is no coverage in the literature of how this structural change in union composition has affected the Mexican labour market. There is also a gap in the literature in studying the effects of unionisation in Mexico as there only exist a handful of papers (Davis and Coleman, 1989; Cragg and Epelbaum, 1996; Fairris, 2003; Fairris and Levine, 2004; Fairris, 2006; Lévesque and Murray, 2005) which analyse the effects of unions.

The present study has some caveats associated with it which must be taken into consideration. The selection equation which we specified may be improved through the inclusion of further identifying instruments. A different specification might result in more selection effects being found. This may not necessarily be possible with the data set which we utilised, therefore we suggest exploring alternatives such as the *Encuesta Nacional de Empleo* (ENE) or the more recent and redesigned *Encuesta Nacional de Ocupacion y Empleo* (ENOE).

The wage curve findings we report may be sensitive to the application instrumental variables technique. Blanchflower and Oswald (2005) suggest the use of lags of the local rate of unemployment as an identifying instrument. This procedure may change the result of the rejection of the the existence of the wage curve. Furthermore the wage curve they estimate is based on industry level rate of unemployment, and this relationship may not hold for state level unemployment rates.

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A Full results

The dependent variable of interest LABSTAT was coded on the basis of five mutually exclusive realisations in the labour market (*viz*, "Out of the labour force", "Unemployed", "Informal", "Formal Private" and "Formal Public"). The specification estimated for both 2000 and 2003 comprised the exogenous variables in our Mincerian earnings equation: age and its quadratic, marital status, educational qualifications, and the local unemployment rate.

Our set of identifying instruments were comprised of variables which are thought to shift participation in a particular employment sector but not wages. The most appropriate instruments that could be obtained from the data were variables describing the household of the respondent. These are comprised of a variable controlling for the household head, the log of the minimum wage, four mutually exclusive variables capturing the number of children under 12 years in the respondent's household, a continuous variable for the number of individuals over 12 years in the respondent's household, as well as a set of controls for the birthplace of the respondent.

One of the important considerations that must be taken into account is that our set of instruments must be relevant and exogenous. If this is not the case then we would be faced with a problem of identification, as they do not only have an effect upon participation but also wage determination. In order to put these fears to rest we devised a crude test by introducing these into an austere specification of the wage earnings equation. In both instances the joint statistical significance of the instruments was rejected.

Table A1 Multinomial Model Estimates

P(Labstat=1) OLF			P(Labstat=2)Informal		P(Labstat=3)			
	2000	2003	2000	2003	2000	2003	2000	2003
Constant	12.0431***	17.0188*	-6.7151*	-12.6030	-12.3081***	-11.4425	-15.3357***	-25.9203**
	(4.1918)	(9.0168)	(4.0268)	(8.3855)	(4.0293)	(8.4747)	(4.5228)	(10.0854)
Age	-0.4815***	-0.5025***	0.0621***	0.0489***	0.1092***	0.12***	0.2905***	0.3156***
	(0.0153)	(0.0138)	(0.0148)	(0.0130)	(0.0149)	(0.0133)	(0.0182)	(0.0187)
Age^2	0.0065***	0.0066***	-0.0006***	-0.0005***	-0.0015***	-0.0017***	-0.003***	-0.0032***
	(0.0002)	(0.0002)	(0.0002)	(0.0002)	(0.0002)	(0.0002)	(0.0002)	(0.0002)
Married	-0.3326***	-0.3978***	0.3079***	0.2441***	0.7132***	0.6269***	0.5603***	0.5537***
	(0.0904)	(0.0822)	(0.0833)	(0.0720)	(0.0833)	(0.0726)	(0.0931)	(0.0881)
Household	-0.086	-0.2694***	0.7476***	0.678***	0.9296***	0.8335***	1.0047***	0.823***
Head	(0.1003)	(0.0904)	(0.0931)	(0.0800)	(0.0932)	(0.0808	(0.1071)	(0.1015)
Children	-0.2517***	-0.2085***	0.2434***	0.1835***	0.1607***	0.1713***	0.146**	0.0529
1-2	(0.0622)	(0.0575)	(0.0604)	(0.0546)	(0.0604)	(0.0551)	(0.0682)	(0.0670)
Children	-0.6586***	-0.4323***	0.3262***	0.3936***	0.2139*	0.2547**	0.1614	0.0345
3-5	(0.1220)	(0.1242)	(0.1109)	(0.1072)	(0.1110)	(0.1082)	(0.1236)	(0.1327)
Children	-0.5513	-0.6049	1.0095	0.1426	0.7650	-0.025	0.5335	0.4678
6+	(0.7732)	(0.5991)	(0.7210)	(0.5275)	(0.7249)	(0.5427)	(0.8848)	(0.7334)
Primary	-1.0836***	-1.2935***	-0.2451	-0.5016***	0.3712**	-0.0499	1.1461***	0.7168**
	(0.1654)	(0.1717)	(0.1600)	(0.1632)	(0.1633)	(0.1690)	(0.2748)	(0.3379)
Secondary	-0.437***	-0.611***	-0.7173***	-0.9742***	0.4373**	0.1327	2.6239***	2.3188***
	(0.1716)	(0.1775)	(0.1670)	(0.1702)	(0.1699)	(0.1754)	(0.2776)	(0.3390)
Preparatory	-0.5252***	-0.8851***	-0.7482***	-1.1106***	0.4624**	0.0640	2.4404***	1.9704***
	(0.1710)	(0.1746)	(0.1661)	(0.1667)	(0.1690)	(0.1720)	(0.2777)	(0.3374)
University	-0.1637	-0.2643	-0.9333***	-1.1953***	0.1387	-0.1614	3.1221***	2.9148***
	(0.1626)	(0.1692)	(0.1580)	(0.1621)(0.1612)	(0.1676)	(0.2703)	(0.3319)	
ln(min wage)	-0.2626	-0.9064	1.1847**	2.0557*	1.771***	1.5490	0.9189	2.3334*
	(0.5993)	(1.2648)	(0.5756)	(1.1763)	(0.5759)	(1.1888)	(0.6453)	(1.4139)
hhmembers	0.0368*	0.0281	-0.0067	0.0142	0.0146	0.0214	0.0254	0.0359*
	(0.0188)	(0.0178)	(0.0182)	(0.0168)	(0.0183)	(0.0171)	(0.0210)	(0.0213)
Local Unemp.	-0.1403**	-0.1166***	-0.0706	-0.1484***	-0.1534***	-0.1018***	-0.0264	-0.1699***
	(0.0578)	(0.0310)	(0.0554)	(0.0288)	(0.0554)	-0.0290	(0.0604)	(0.0336)
Small–Hsiao	176.55**	148.42	256.49***	163.98*	223.457***	171.09**	186.63***	143.72
Birthplace Controls	Yes	Yes						
Pseudo–R^2	0.1920	0.1989						
N	76631	54016						

Notes: Unemployment is the base outcome, due to the theil normalisation estimates for this category are not obtained
***,**,* denote significance at 1, 5 and 10 per cent respectively.
Standard Errors reported in parentheses.
Small–Hsiao test reported is for exclusion of category under which it appears

Table A2 OLS Estimates for Mincerian Earnings Equation for Mexico, 2000 and 2003

	2000				2003			
	Pooled	Informal	Private	Public	Pooled	Informal	Private	Public
Constant	-0.592***	-0.377***	-0.124	-0.185	1.368***	1.670***	1.672***	2.321***
	(.0357)	(.0769)	(.0933)	(.3359)	(.0435)	(.0866)	(.1280)	(.4560)
Age	0.042***	0.034***	0.032***	0.030***	0.039***	0.028***	0.034***	0.006
	(.0015)	(.0030)	(.0030)	(.0088)	(.0018)	(.0034)	(.0041)	(.0120)
Age^2	-0.00044***	-0.00037***	-0.00027***	-0.00024**	-0.00042***	-0.00031***	-0.00030***	0.00004
	(.00002)	(.00004)	(.00004)	(.00009)	(.00002)	(.00004)	(.00006)	(.00013)
Married	0.126***	0.159***	0.046***	0.097***	0.128***	0.138***	0.073***	0.107***
	(.0061)	(.0103)	(.0111)	(.0202)	(.0074)	(.0115)	(.0144)	(.0251)
Primary	0.183***	0.198***	0.067**	0.178	0.113***	0.106***	0.067**	0.231*
	(.0135)	(.0176)	(.0222)	(.1328)	(.0160)	(.0203)	(.0263)	(.1341)
Secondary	0.366***	0.371***	0.233***	0.583***	0.291***	0.300***	02.40***	0.578***
	(.0151)	(.0264)	(.0244)	(.1409)	(.0182)	(.0300)	(.0299)	(.1520)
Preparatory	0.437***	0.456***	0.307***	0.498***	0.330***	0.347***	0.287***	0.476***
	(.0152)	(.0262)	(.0250)	(.1395)	(.0178)	(.0288)	(.0306)	(.1486)
University	0.703***	0.671***	0.628***	0.888***	0.583***	0.546***	0.593***	0.865***
	(.0149)	(.0276)	(.0231)	(.1461)	(.0178)	(.0313)	(.0280)	(.1638)
Temp. Contract	-0.060**	0.055	-0.118***	-0.214	-0.045	0.114**	-0.133***	-0.356***
<2 months	(.0205)	(.0454)	(.0208)	(.2034)	(.0285)	(.0565)	(.0340)	(.1016)
Temp. Contract	-0.109***	-0.103**	-0.109***	-0.208***	-0.068***	-0.045	-0.058**	-0.257***
2–6 months	(.0160)	(.0342)	(.0186)	(.0557)	(.0191)	(.0422)	(.0216)	(.0629)
Temp. Contract	-0.090***	-0.038	-0.097***	-0.153	-0.198***	-0.223	-0.142***	0.367**
6> months	(.0143)	(.0345)	(.0162)	(.0440)	(.0179)	(.0410)	(.0199)	(.0624)
Verbal Contract	-0.236***	-0.262***	-0.175***	-0.153	-0.198***	-0.223	-0.142***	0.367**
	(.0077)	(.0098)	(.0138)	(.1279)	(.0093)	(.0113)	(.0199)	(.1747)
Local Unemp.	0.062***	0.078***	0.051***	0.024	0.120***	0.150***	0.093***	0.079**
	(.0052)	(.0082)	(.0077)	(.0150)	(.0092)	(.0147)	(.0128)	(.0297)
Informal	0.017**	–	–	–	0.00031	–	–	–
	(.0065)				(.0079)			
Public	0.178***	–	–	–	0.213***	–	–	–
	(.0101)				(.0127)			
SCT	–	-0.037	-0.201***	0.237	–	-0.101**	-0.163***	-0.063
		(.0369)	(.0291)	(.0477)		(.0388)	(.0374)	(.0635)
State Controls	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Ind. Controls	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
R^2	0.4046	0.3423	0.4006	0.3922	0.3955	0.3201	0.3863	0.3862
N	59547	26131	28117	5299	39605	18798	17299	3508

Notes: ***,**,* denote significance at 1, 5 and 10 per cent respectively.
The standard errors reported in parentheses were calculated using the White(1980) sandwich estimator
The pay measure is net of taxes and is measured as the log of the hourly wage in constant 1990 prices
The regressions included 12 industry controls, and 31 regional controls

Table A3 Quantile Regression Estimates for Augmented Mincerian Earnings Equation at the median, 2000 and 2003

	2000				2003			
	Pooled	Informal	Private	Public	Pooled	Informal	Private	Public
Constant	-0.551***	-0.380***	-0.024	0.024	1.462***	1.722***	1.809***	2.106***
	(.0345)	(.0588)	(.0986)	(.2048)	(.0408)	(.0785)	(.1282)	(.1282)
Age	0.041***	0.037***	0.028***	0.025***	0.037***	0.027***	0.030***	0.007
	(.0014)	(.0024)	(.0032)	(.0065)	(.0016)	(.0029)	(.0041)	(.0097)
Age^2	-0.0004***	-0.0004***	-0.0002***	-0.0002**	-0.0004***	-0.0003***	-0.0003***	.00004
	(.00002)	(.00003)	(.00004)	(.00008)	(.00002)	(.00004)	(.00006)	(.0001)
Married	0.128***	0.156***	0.043***	0.066***	0.113***	0.127***	0.056***	0.085***
	(.0059)	(.0107)	(.0118)	(.0219)	(.0068)	(.0102)	(.0142)	(.0314)
Primary	0.173***	0.169***	0.071***	0.321**	0.128***	0.123***	0.061**	0.286
	(.0128)	(.0189)	(.0257)	(.1361)	(.0157)	(.0192)	(.0307)	(.2022)
Secondary	0.329***	0.292***	0.224***	0.629***	0.256***	0.254***	0.192***	0.648***
	(.0142)	(.0227)	(.0279)	(0.1361)	(.0175)	(.0267)	(.0339)	(.2038)
Preparatory	0.395***	0.356***	0.278***	0.562***	0.302***	0.302***	0.239***	0.543***
	(.0142)	(.0225)	(.0283)	(.1361)	(.0170)	(.0257)	(.0342)	(.2032)
University	0.675***	0.593***	0.618***	0.973***	0.548***	0.495***	0.535***	0.952***
	(.0137)	(.0212)	(.0261)	(.1350)	(.0166)	(.0274)	(.0317)	(.2020)
Temp. Contract	-0.054***	0.075*	-0.100***	-0.067	-0.052**	0.165***	-0.109***	-0.351
<2 months	(.0201)	(.0434)	(.0249)	(.1721)	(.0251)	(.0438)	(.0309)	(.2231)
Temp. Contract	-0.079***	-0.116***	-.080***	-0.209***	-0.082***	-0.041	-0.086***	-0.307***
2–6 months	(.0165)	(.0376)	(.0205)	(.0760)	(.0183)	(.0342)	(.0228)	(.0943)
Temp. Contract	-0.073***	-0.085**	-0.070***	-0.145***	-0.033**	0.015	-0.053***	-0.102
6> months	(.0140)	(0.0341)	(.0177)	(.04474)	(.0159)	(.0310)	(.0192)	(.0830)
Verbal Contract	-0.233***	-0.266	-0.143***	-0.311**	-0.192***	-0.214***	-0.139***	0.224
	(.0072)	(.0107)	(.0153)	(.1461)	(.0084)	(.0101)	(.0212)	(.2447)
Local Unemp.	(.2447)	0.073***	0.049***	0.007	0.106***	0.147***	0.077***	0.055*
	(.0050)	(.0088)	(.0080)	(.0153)	(.0084)	(.0129)	(.0128)	(.0337)
Informal	0.012**	–	–	–	0.0003	–	–	–
	(.0058)				(.0069)			
Public	0.176***	–	–	–	0.215***	–	–	–
	(.0099)				(.0118)			
SCT	ψ	ψ	-0.231***	ψ	ψ	-0.086**	-0.162***	ψ
			(.0305)			(.0341)	(.0371)	
State Controls	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Ind. Controls	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Pseudo–R^2	0.2587	0.2151	0.2428	0.2503	0.2484	0.1927	0.2344	0.2514
N	59547	26131	28117	5299	39605	18798	17299	3508

Notes: ***,**,* denote significance at 1, 5 and 10 per cent respectively.
ψ denotes selection correction terms not included in regression
The pay measure is net of taxes and is measured as the log of the hourly wage in constant 1990 prices
The regressions included 12 industry controls, and 31 regional controls

Table A4 Quantile Regression Estimates for Augmented Mincerian Earnings Equation at the 10th quantile, 2000 and 2003

	2000				2003			
	Pooled	Informal	Private	Public	Pooled	Informal	Private	Public
Constant	-0.727***	-0.801***	-0.314***	-0.420*	1.106***	1.070***	1.391***	1.562***
	(.0476)	(.1061)	(.1217)	(.2483)	(.0668)	(.0999)	(.1657)	(.3994)
Age	0.033***	0.029***	0.022***	0.020**	0.033***	0.0276***	0.026***	0.018
	(.0137)	(.0041)	(.0039)	(.0087)	(.0027)	(.0040)	(.0053)	(.0127)
Age^2	-0.0004***	-0.0004***	-0.0002***	-0.0002*	-0.0004***	-0.0004***	-0.0003***	-0.0001
	(.00002)	(.00005)	(0.00005)	(.0001)	(.00003)	(.00005)	(.00007)	(.0001)
Married	0.104***	0.160***	0.032**	0.111***	0.119***	0.155***	0.061***	0.077
	(.0079)	(.0142)	(.0144)	(.0287)	(.0110)	(.0172)	(.0186)	(.0402)
Primary	0.118***	0.143***	.064**	0.1353	0.077***	0.058*	0.055	0.090
	(.0175)	(.0248)	(.0317)	(.1395)	(.0255)	(.0332)	(.0401)	(.2780)
Secondary	0.234***	0.279***	0.159***	0.515***	0.191***	0.119***	0.168***	0.301
	(.0196)	(.0355)	(.0344)	(.1395)	(.0285)	(.0397)	(.0440)	(.2834)
Preparatory	0.271***	0.328***	0.204***	0.423***	0.215***	0.193***	0.180***	0.268
	(.0195)	(.0350)	(.0347)	(.1386)	(.0277)	(.0383)	(.0443)	(.2830)
University	0.421***	0.456***	0.370***	0.685***	0.351***	0.276***	0.348***	0.5188*
	(.0191)	(.0361)	(.0324)	(.1370)	(.0276)	(.0376)	(.0414)	(.2813)
Temp. Contract	-0.073***	0.047	-0.095***	-0.447**	-0.125***	-0.045	-0.192***	-0.193*
<2 months	(.0273)	(.0564)	(.0312)	(.2258)	(.0404)	(.0740)	(.0410)	(.1068)
Temp. Contract	-0.053**	0.0002	-0.075***	-0.080	-0.044	-0.018	-0.060**	-0.084
2–6 months	(.0224)	(.0481)	(.0255)	(.0909)	(.0293)	(.0587)	(.0295)	(.1154)
Temp. Contract	-0.061***	0.062	-0.060***	-0.185***	-0.011	0.051	-.028	-0.049
6> months	(.0190)	(.0437)	(.0222)	(.0597)	(.0253)	(.0522)	(.0252)	(.1089)
Verbal Contract	-0.139***	-0.141***	-0.177***	-0.187	-0.021***	-0.125***	-0.145***	0.153
	(.0099)	(.0139)	(.0195)	-0.1604	(.0138)	(.0174)	(.0284)	(.1186)
Local Unemp.	0.053***	0.071***	0.020**	0.024	0.115***	0.115***	0.109***	0.010
	(.0067)	(.0113)	(.0100)	(.0196)	(.0134)	(.0220)	(.0167)	(.0448)
Informal	-0.153***	–	–	–	-0.166***	–	–	–
	(.2539)				(.0111)			
Public	0.254***	–	–	–	0.251***	–	–	–
	(.0137)				(.0197)			
SCT	ψ	-0.107**	-0.147***	ψ	ψ	ψ	-0.125***	ψ
		(.0471)	(.0368)				(.0478)	
State Controls	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Ind. Controls	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Pseudo–R^2								
N	59547	26131	28117	5299	39605	18798	17299	3508

Notes: ***,**,* denote significance at 1, 5 and 10 per cent respectively.
ψ denotes selection correction terms not included in regression
The pay measure is net of taxes and is measured as the log of the hourly wage in constant 1990 prices
The regressions included 12 industry controls, and 31 regional controls

Table A5 Quantile Regression Estimates for Augmented Mincerian Earnings Equation at the 90th quantile, 2000 and 2003

	2000				2003			
	Pooled	Informal	Private	Public	Pooled	Informal	Private	Public
Constant	-0.180***	0.379***	0.309*	0.758**	1.849***	2.250***	2.296***	1.882***
	(.0683)	(.0990)	(.1774)	(.3099)	(.0878)	(.1048)	(.2401)	(.4594)
Age	0.045***	0.029***	0.036***	0.021**	0.041***	0.028***	0.034***	0.021
	(.0028)	(.0040)	(.0057)	(.0097)	(.0035)	(.0041)	(.0075)	(.0140)
Age^2	-0.0004***	-0.0003***	-0.0002***	-0.00009	-0.0004***	-0.0003***	-0.0002**	-0.0001
	(.00004)	(.00005)	(.00008)	(.0001)	(.00004)	(.00005)	(.0001)	(.0001)
Married	0.112***	0.138***	0.015	0.062*	0.121***	0.103***	0.071***	0.130***
	(.0116)	(.0178)	(.0209)	(.0337)	(.0144)	(.0183)	(.0255)	(.0453)
Primary	0.168***	0.178***	0.040	-0.110	0.092***	0.108***	0.037	0.501*
	(.0257)	(.0318)	(.0451)	(.2152)	(.0338)	(.0351)	(.0546)	(.2915)
Secondary	0.414***	0.393***	0.268***	0.396*	0.377***	0.350***	0.286***	0.972***
	(.0288)	(.0384)	(.0492)	(.2158)	(.0375)	(.0414)	(.0607)	(.2927)
Preparatory	0.516***	0.512***	0.388***	.263	0.401***	0.374***	0.355***	0.835***
	(.0288)	(.0382)	(.0498)	(.2155)	(.0370)	(.0402)	(.0617)	(.2922)
University	0.891***	0.795***	0.862***	0.715***	0.792***	0.647***	0.838***	1.340***
	(.0276)	(.0359)	(.0458)	(.2149)	(.0359)	(.0390)	(.0562)	(.2908)
Temp. Contract	-0.062	0.094	-0.152***	-0.439*	-0.010	0.123	-0.112**	-0.510
<2 months	(.0399)	(.0724)	(.0439)	(.2589)	(.0535)	(.0795)	(.0549)	(.3445)
Temp. Contract	-0.188***	-0.186***	-0.163***	-0.463***	-0.085**	-0.101	-0.047	-0.335**
2–6 months	(.0329)	(.0628)	(.0362)	(.1118)	(.0391)	(.0626)	(.0403)	(.1366)
Temp. Contract	-0.124***	-0.038	-0.109***	-0.150**	-0.043	0.105*	-0.066**	-0.161
6> months	(.0279)	(.0581)	(.0315)	(.0688)	(.0339)	(.0570)	(.0333)	(.1223)
Verbal Contract	-0.299***	-0.407***	-0.107***	-0.096	-0.290***	-0.378***	-0.122**	0.188
	(.0146)	(.0165)	(.0265)	(.1862)	(.0182)	(.0182)	(.0378)	(.1229)
Local Unemp.	0.072***	0.099***	0.056***	0.038*	0.136***	0.160***	0.082***	0.211***
	(.0099)	(.0149)	(.0141)	(.0230)	(.0179)	(.0233)	(.0230)	(.0450)
Informal	0.165***	–	–	–	0.157***	–	–	–
	(.0123)				(.0252)			
Public	0.120***	–	–	–	0.162***	–	–	–
	(.0194)				(.0252)			
SCT	ψ	ψ	-0.279***	ψ	ψ	ψ	-0.214***	
			.(0555)				(.0691)	
State Controls	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Ind. Controls	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Pseudo–R^2								
N	59547	26131	28117	5299	39605	18798	17299	3508

Notes: ***,**,* denote significance at 1, 5 and 10 per cent respectively.
ψ denotes selection correction terms not included in regression
The pay measure is net of taxes and is measured as the log of the hourly wage in constant 1990 prices
The regressions included 12 industry controls, and 31 regional controls